Tax Incidence when Quality Matters: Evidence from the Beer Market

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Abstract

What are the effects of a per-unit tax on the sales of a good? The answer in standard economics textbooks is a reduction in quantity and an increase in price. However, the homogeneous-product assumption in such treatments is not always correct because the taxed product is frequently sold under numerous differentiated brands. In general, quantity and price adjustments to a per-unit tax increase may vary differentially by brand, which implies that tax incidence requires a more thorough analysis than the standard one. In this paper we test a popular theoretical prediction that the effects of a per-unit tax on sales may be different for low quality brands than for high quality brands. We exploit two natural experiments in the beer industry that capture exogenous variations on per-unit charges (excise taxes and shipping costs) to study if quality is an important determinant of equilibrium quantity, prices and advertising expenditures. We find that, when faced with higher per-unit charges, the equilibrium price and quantity of a brand are usually positively associated with its quality level; similarly, higher per-unit charges appear to provide firms with an incentive to shift advertising effort towards brands of higher quality. We discuss implications for public policy and firm profitability.

JEL codes: L13, H22, L66

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1. Introduction

What are the effects of a per-unit tax on the consumption of a good? The answer in standard economics textbooks is a reduction in quantity traded and an increase in price. However, the homogeneous product assumption in such treatments is not always correct because the taxed product is often sold under numerous brand names, each being perceived differently by consumers, for example with varying quality levels. As a consequence, quantity and price adjustments to the tax increase may vary differentially by brand, which implies that tax incidence requires a more thorough analysis than the standard one. Specifically, a per-unit tax is imposed on the *quantity* sold, which may make consumption and production of certain (untaxed) product attributes relatively cheaper to purchase and manufacture. For example, a tax increase of $0.50 per gallon on all gasoline sales may induce consumers to substitute towards higher grade gasoline as quality becomes relatively less expensive after the tax increase.[1]

The concept of how a per-unit charge may differentially affect products with different attributes was introduced by Alchian and Allen (1964) and Barzel (1976), albeit from different angles. A key assumption in both of these approaches is that consumers and producers, all else equal, prefer a higher level of the untaxed characteristic. To maintain our analysis consistent with this assumption, in what follows we refer to a single untaxed product characteristic that meets this criterion: quality.[2] Alchian and Allen assume supply remains constant and study consumer incentives when the same per-unit fee (i.e. transportation cost) needs to be paid to purchase either a low quality (low-priced) good or a high quality (high-priced) good. Since consumers value quality, the high-priced good becomes relatively cheaper to acquire and hence its demand

[1] In contrast, an ad valorem tax is applied to the price of a product thereby implicitly taxing all characteristics.
[2] While products differ across different dimensions (e.g. color, size, etc.), it is not clear which of these characteristics would fit this criterion. A dimension that also fits this *vertical* nature of product differentiation is durability, an example that Barzel explored in his work but that does apply to the product we study. Our model below accommodates for both horizontal and vertical product differentiation.

relative to low quality products increases. In principle, the same intuition could be applied to the case when a per-unit tax raises the price of high and low quality goods by the same amount.[3]

Barzel analyzes the issue from a supply and demand perspective in a perfectly competitive market. In equilibrium, the marginal cost of producing each product characteristic (taxed and untaxed) is equal to society's marginal valuation of the characteristic, but when a per-unit tax is imposed, the equilibrium is distorted and adjustments need to be made. While Barzel analysis focuses on the firm's incentive to modify the product composition by upgrading quality to restore the equilibrium and minimize the impact of the per-unit tax, such adjustment also comes from consumers' incentives to shift their purchase decisions towards higher quality products (p. 1181). In principle, any per-unit charge, such as a shipping cost, has similar implications as a per-unit tax; in this paper we consider *two* different types of per-unit charges to study the effects predicted by Alchien-Allen and Barzel: per-unit taxes and shipping costs.

While Alchian-Allen and Barzel address the problem from different angles, they generate similar predictions, two of which are the focus of this study.[4] The first prediction is that the amount of quality traded in the market should increase when there is an increase in the per-unit charge. This prediction implies that a per-unit charge should reduce quantity traded by a smaller amount for high quality products than for low quality products (we will often refer to this prediction as the "quantity effect"); the reason for this is that the negative impact that a per-unit charge has on sales will be lessened for high quality products because of consumers' shift towards higher quality. The second prediction is related to equilibrium price: the demand shift

[3] However, as we will show, in differentiated products markets a per-unit tax has complex implications on equilibrium prices, so such uniform price increases (and the effects predicted by this model) are not guaranteed.
[4] Barzel also discusses how per-unit taxes and ad-valorem taxes have different effects on sales (when firms can modify the quality content of their products). We do not address this prediction as we assume that quality remains fixed and because of concerns with the endogeneity of ad-valorem taxes (see section 4.2 for a discussion on this issue).

towards higher quality products implies that equilibrium price will be relatively higher for high quality products than for low quality products (we will often refer to this prediction as the "price effect").[5]

Empirical research has focused on testing the first prediction. Bertonazzi, Maloney, McCormick (1993) and Hummels and Skiba (2004) provide evidence that longer distances between consumers and the product location are associated with a larger demand for higher quality goods. Sobel and Garret (1997) find that the market share of premium (i.e. high quality) cigarettes increases in the presence of larger per-unit taxes. Barzel, Johnson (1978) and Sumner and Ward (1981) conduct a "mixed test" of the two predictions, namely whether the quantity-weighted average price (across all brands) increases by more than the per-unit tax; however, these studies are inconsistent in their findings. A general shortcoming of previous work is that analyses have been limited by the absence of product quality measures or by the unavailability of detailed brand-level data. As a result, conclusions have been based in terms of *average* prices (or price indices) or coarser measures of quality (e.g. premium vs. generic cigarettes).[6] More importantly, the questions of interest have not been studied in a more realistic model that considers the simultaneous demand and supply shifts caused by per-unit taxes.

While predictions by Alchian-Allen and Barzel are appealing and have received some empirical support, theoretical work has shown that these predicted effects may not always hold. Leffler (1982), for example, considers a model where the amount of quality consumed is

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[5] This is an effect that is only explicitly recognized by Barzel, albeit using a supply side story: because quality is costly to produce, the firm's upgrade in quality generates an equilibrium price that is higher than the equilibrium price when quality is constant; however (and consistent with our assumption below that quality is constant in the period of study) this conclusion can also be reached using a demand side story: given two brands with different quality levels, some consumers will shift their demand from the low quality brand to the high quality brand. Conversely, Alchian and Allen assume that all prices increase by the same amount, but this assumption is only met in the extreme case of perfect competition.

[6] Other related empirical literature includes Feenstra (1988) finds that a quota restriction on Japanaese car imports into the US incentivizes consumers to upgrade the quality of cars imported from Japan.

measured in relative terms (with respect to quantity) and finds that the effect of a per-unit tax on product quality is ambiguous. Also, substitution towards higher quality goods may fail to hold under certain conditions (Gould and Segall, 1968; Borcherding and Silberberg, 1978). Further, we show that Alchian-Allen and Barzel predictions are indeterminate in a partial equilibrium model of imperfect competition that incorporates market features not previously studied.

The intuitive but ambiguous effects predicted by these two theories make empirical evidence particularly interesting to investigate. We make use of two experiments in the beer industry that capture exogenous variations on per-unit charges. The first experiment is the one-time 100% increase in the federal excise tax on beer sales that took place in January of 1991. This increase (from $9 to $18 per barrel) was imposed on all beer sales (domestic and imports) and was the largest federal excise tax increase in the history of the US beer industry. The second experiment makes use of the high transportation costs of beer that results from substantial shipping distances and the heaviness of the product. Importantly, our data set contains a large variation in shipping distances across cities as well as across brewers. Since a shipping cost is a per-unit charge that is mathematically equivalent to a per-unit tax,[7] we use shipping distances to proxy for shipping costs.

We exploit these two experiments to study whether equilibrium quantity and prices vary differentially according to each brand's quality level when per-unit charges increase. An intuitive extension of the effects that we are interested in is that a higher per-unit charge may provide firms with an incentive to increase the advertising intensity for high quality products with respect to low quality products. Thus, we also study the effects of larger per-unit charges on advertising intensities of brands with different quality levels.

[7] We assume that the per-mile shipping cost of a unit of volume is constant. The plausibility of this assumption is explored in section 3.

A main contribution of our work is the computation of a quality estimate for each of the numerous brands in a market for differentiated products. Our approach in overcoming this hurdle is to approximate a brand's quality level with an estimate of its brand equity.[8] We adopt a commonly used definition of brand equity: the valuation consumers (and consequently firms) give to a particular brand.[9] As is common in the economics and marketing literatures (e.g. Berry, 1994; Goldfarb, Lu, Moorthy, 2007; Sriram, Balachander and Kalwani, 2007) we operationalize the measurement of brand equity by estimating demand for the numerous brands in this market. Specifically, brand equity is defined by the time-invariant brand fixed-effect that shifts the demand of the differentiated product.[10]

Importantly, our structural approach allows us to separate the brand equity effect on sales from those effects given by prices, advertising, and other variables. Specifically, our measure of brand equity takes into account: a) the strategic nature of firms' behavior (by accounting for the endogeneity of prices and advertising), and b) consumers' complex substitution preferences. Another advantage of our method is that instead of using potentially subjective expert rankings of product quality, we choose (loosely speaking) to ask this question directly from the data. The brand equity estimates that we obtain seem very plausible, giving us confidence about the validity of our approach. For example, the best selling beers as well as imports and super-

[8] Another possibility is to use a hedonic approach to obtain a measure of quality that is a function of observed product characteristics as in, for example, Feenstra (1988). We are hesitant to follow this approach, however, as specific product attributes are more difficult to measure in this market than in other industries (e.g. automobiles). Another possible drawback from hedonic models is that theory does not provide a straightforward choice of functional form.

[9] Brand equity should not be confused with brand value. The former is a measure of the premium that consumers place on the consumption of a particular brand, whereas the latter is defined as the monetary value that brand equity represents for the manufacturing firm. Brand equity and brand value are intimately related, however. For a review of the literature see Ailawadi, Lehmann and Neslin (2003).

[10] In the economics literature the concept of brand equity is (usually) not explicitly recognized as such. Rather, it was originally introduced as an econometric problem by Berry (1994) who suggested that the existence of an unobserved product characteristic (such as brand loyalty or quality) can be a source of endogeneity in the demand equation. This econometric problem has been (partially) addressed and captured via the introduction of brand fixed effects in several applications (e.g. Nevo, 2001); a by-product of this approach is that the computed fixed effects can be used as an estimate of brand equity. Section 3 provides further details on this issue.

premium brands have the largest brand equity estimates, while budget and less popular brands have the lowest brand equity estimates (see table 3 for a complete ranking of brand equity estimates and section 5.2 for further discussion).

Our treatment of quality is not without caveats, however. We assume that quality for a given brand during the period of study is constant. But as a result of the federal tax change, firms may have an incentive to modify the quality levels of their brands, as originally suggested by Barzel. Thus, our price and quantity tests do not capture this supply side effect.[11] We carry out some robustness checks in our empirical analysis to reduce concerns about this assumption, but acknowledge at the outset that the fixed quality assumption is a limitation.

A secondary contribution of our work is to construct a theoretical model that incorporates key features of imperfectly competitive markets: product differentiation, multiproduct firms and the strategic nature of price and advertising. The main objective of this model is to show that the effects predicted by Alchian-Allen and Barzel are ambiguous when complex features of a market are incorporated, thereby making the need for empirical evidence even more pressing.

Our results indicate that, despite their ambiguity, the price and quantity effects of interest are largely present in the beer market. In particular, we find that price and quantity effects emerge when the difference in quality between a pair of brands is large enough. The quantity effect is the least common: it is present in approximately 40% of all possible pairs of brands. When compared to the case of homogeneous quality, this finding implies higher tax revenues but a less effective reduction in alcohol consumption. The price effect is more prevalent than the quantity effect: it is present in 78% of the cases, which, together with the quantity effect, implies

[11] One way to address this issue is by making quality a choice variable. Recent attempts tackle the issue of modeling product varieties (see for example Draganska, Mazzeo and Seim, 2007 and references therein); but in this literature product choices are easily observed (e.g. number of flavors) and not estimated thereby making these techniques impractical in our case.

that the negative impact of the tax increase on a firm's profit decreases with product quality. Finally, we find that our hypothesized advertising effect is the largest and most prevalent: it is present in 85% of the cases.

Section 2 presents the theoretical model. Section 3 describes the data and section 4 provides details of the empirical strategy. The results are presented in section 5, and conclusions are given in section 6.

2. Theoretical Model

The purpose of this section is to show the ambiguous effects of an increase in the per-unit tax on relative quantities, prices and advertising, between low and high quality brands. Let's denote the price, quantity, advertising and quality of a brand *j* as p_j, q_j, A_j, and ξ_j respectively, and the per-unit charge (either a per-unit tax or a per-unit shipping cost, or a combination of both) as *T*. If the quantity effect is present, the condition $\partial(q_j / q_k)/\partial T > 0$ should hold for any $\xi_j > \xi_k$. Similarly, if the price effect is present, the condition $\partial(p_j / p_k)/\partial T > 0$ should also hold. Finally, if our conjecture that higher per-unit charges increase relative advertising levels is true, the condition $\partial(A_j / A_k)/\partial T > 0$ should be observed. After some manipulation, these conditions can be written as:

(1) $$\frac{1}{z_k^2}\left[z_k \frac{\partial z_j}{\partial T} - z_j \frac{\partial z_k}{\partial T} \right] > 0, \text{ where } z \in \{q, p, A\}$$

Whether inequality (1) holds depends on: a) how responsive are equilibrium quantity, price and advertising to a tax change ($\partial z / \partial T$), and b) the levels of z_j and z_k. If the levels of z_j and z_k are taken as given, one could analyze the plausibility of inequality (1) by studying the magnitude of the derivative terms ($\partial z / \partial T$); to do this we next consider a partial equilibrium model.

We assume that firms compete in prices and advertising. While firms also determine products' quality levels, we assume that for the period of study product quality remains constant. This assumption is used to simplify the exposition of our model, but there is a practical reason for this assumption: it allows us to econometrically identify quality via the inclusion of brand fixed effects in demand estimation (see section 3 for details). Demand for product *j* is a function of all prices ($p = p_1,..., p_J$), advertising expenditures ($A = A_1,..., A_J$) and quality levels ($\xi = \xi_1,...,\xi_J$) in the market: $q_j(p, A, \xi)$.[12] Thus, the quantity derivative $\partial q_j / \partial T$ depends on how equilibrium price and advertising levels shift as a result of the tax change (given our assumption of constant quality levels) and is given by:

$$\partial q_j / \partial T = \sum_{k=1}^{J} \frac{\partial q_j}{\partial p_k} \frac{\partial p_k}{\partial T} + \sum_{k=1}^{J} \frac{\partial q_j}{\partial A_k} \frac{\partial A_k}{\partial T}$$

The price derivative $\partial p_k / \partial T$ and advertising derivative $\partial A_k / \partial T$ depend on the strategic nature of firms' behavior and can thus be obtained from firms' first order conditions. Let F_n be the set of brands produced by firm n. Assuming constant marginal costs (c_j) and linear additivity of advertising, the profit of firm n is:

$$\pi_n = \sum_{j \in F_n} (p_j - c_j - T) q_j(p, A, \xi) - \sum_{j \in F_n} A_j$$

Assuming firms compete in Bertrand-Nash fashion[13], product *j*'s first order conditions are:

$$(2) \qquad q_j(p, A, \xi) + \sum_{k \in F_n} (p_k - c_k - T) \frac{\partial q_k}{\partial p_j} = 0, \text{ with respect to } p_j$$

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[12] Section 3 presents details of the specific functional form used to estimate demand.

[13] Rojas (2006) tests different models of competition in the beer industry and finds that firm behavior is explained reasonably well by Bertrand-Nash.

(3) $\sum_{k\in F_n}(p_k - c_k - T)\frac{\partial q_k}{\partial A_j} - 1 = 0$, with respect to A_j

For a given set of demand parameters and marginal costs, expressions (2) and (3) determine the equilibrium prices and the equilibrium advertising levels in the market. Given the complex nature of the model (numerous differentiated products, multiproduct firms and two strategic variables), however, it is not possible to derive explicit formulas for $\partial p_j / \partial T$ and $\partial A_j / \partial T$ so they need to be evaluated via simulation. As a consequence, it is difficult to obtain conditions under which the price effect $\frac{1}{p_k^2}\left[p_k\frac{\partial p_j}{\partial T} - p_j\frac{\partial p_k}{\partial T}\right] > 0$ and our conjecture $\frac{1}{A_k^2}\left[A_k\frac{\partial A_j}{\partial T} - A_j\frac{\partial A_k}{\partial T}\right] > 0$ would hold. Simulations (not shown) indicated that these two inequalities may or may not hold depending on the parameters of the model (marginal costs, demand equation and demand derivatives). This result, in turn, implies that the quantity effect $\frac{1}{q_k^2}\left[q_k\frac{\partial q_j}{\partial T} - q_j\frac{\partial q_k}{\partial T}\right] > 0$ is also theoretically indeterminate.

Intuitively, the indeterminacy of price and quantity effects is in part due to the fact that our analysis is based on a partial equilibrium approach that endogeneizes demand and supply responses (except for quality responses) to different tax levels. In addition, we incorporate several market factors that have not been considered in prior work, and these factors, we believe, contribute to the indeterminacy of price and quantity effects. For example, our model not only captures vertical product differentiation (through the vector of quality parameters ξ)[14], but also horizontal product differentiation (through the demand function via the cross-price

[14] Note that vertical product differentiation is also captured by the magnitude of own-price coefficients in the demand function (a smaller price sensitivity is associated with greater consumer loyalty).

coefficients)[15]. Further, the model allows for multiproduct firms that compete strategically in price and advertising.[16]

One additional implication of the complex forces at play in our model is that prices will not rise by the same amount when there is an increase in a per-unit charge, as assumed by previous empirical tests of the Alchian-Allen effect.[17] In general, our model may be applicable to other industries that are affected by per-unit taxes as it includes features that are frequently observed in many markets.

3. Data

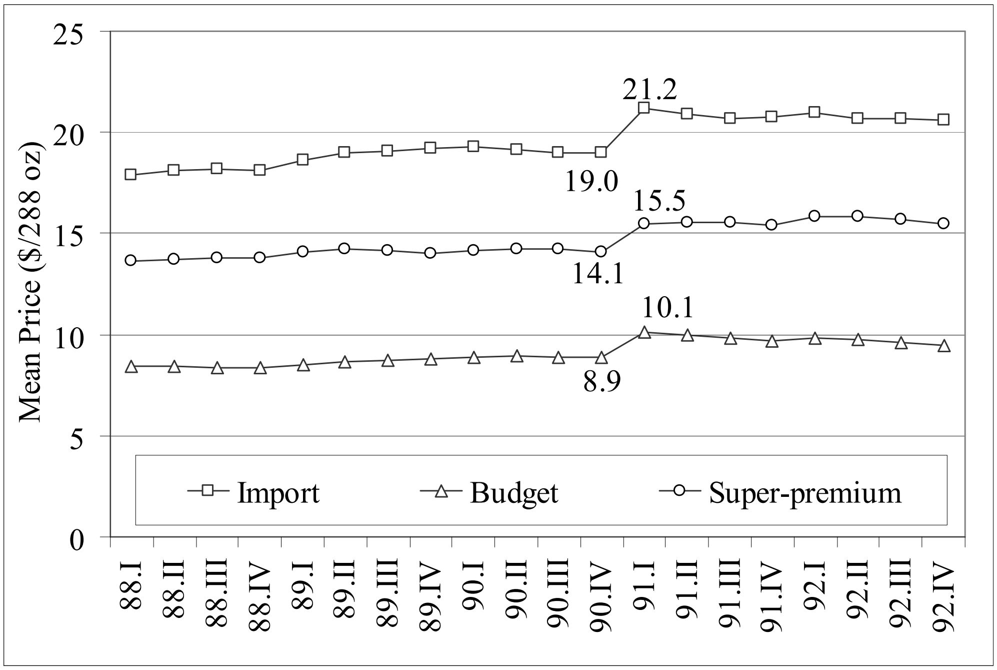
The Federal Excise Tax Increase

In 1990, U.S. Congress approved an increase in the federal excise tax on beer from $9 to $18 per barrel. All brewers and importers were required to pay this tax on all produced units as of January of 1991. This increase, which was equivalent to an additional 65 cents in federal taxes per 288 ounces (a 24-pack), represented the largest federal tax hike for beer in U.S. history.

Figure 1 shows mean quarterly prices (over all cities) for three beer segments using the data set employed in this paper. There is a clear shift in the mean price of all three categories in the first quarter of 1991. All mean increases are higher than the actual tax hike: 220 cents for imports, 140 cents for super-premium beers and 120 cents for budget beers. These mean increases were 238%, 115%, and 85%, respectively, larger than the tax increase.

[15] Our empirical model (section 3) describes the demand model employed, which is consistent with this observation: cross-price coefficients are directly related to the degree of similarity between brands.
[16] While our model is rich as it integrates many key characteristics of the market under study, we do not integrate all of its possible features. For example, we ignore dynamic aspects of firm behavior as well as alternative models of vertical competition. We believe, however, that adding further complexities to the model will keep rendering Allen-Alchian and Barzel effects an empirical matter.
[17] This is not a criticism of the Allen-Alchian concept, it just states that care needs to be taken when assuming that a per-unit charge will increase all brands' prices by the same amount.

Figure 1: Quarterly Mean Prices, Various Beer Segments (1988-1992)



Source: IRI Database, University of Connecticut

The price overshifting pattern shown in figure 1 is also present at the brand level (not shown). Price overshifting was the basis for earlier tests of Barzel's theory: a larger per-unit tax provides firms with an incentive to upgrade quality. If this argument is true here, then the large (within brand) price increases may be reflecting an upgrade in quality in the post-tax-increase period thereby violating our constant quality assumption. But we believe that the explanation for the large price increases is unlikely to be due to an upgrade in quality. First, in our estimation below we directly test whether within brand quality changed upwards in the post-increase period but fail to find evidence for such hypothesis. Second, in prior work (Rojas, 2006) we construct a partial equilibrium model using data prior to the tax increase to simulate the equilibrium prices that would be observed under various modes of competition as a result of the tax increase; all these simulation exercises hold quality constant but almost always predict price overshifting.[18]

[18] These simulations are consistent with the theoretical findings of Anderson, de Palma and Kreider (2001) who show that the strategic interaction in oligopolies with differentiated products may cause an excise tax to be passed on to consumers by more than 100%.

Finally, had a quality upgrade been in place, we would expect to see a gradual and slow change in it, rather than the abrupt shift reflected in the graph.

Table 1: Description and Summary of Statistics of Variables (nominal values)

Variable	Description	Units	Mean	St dev	Min	Max
Price	Average Price per brand	$/288oz	12.10	3.87	0.82	28.97
Quantity	Volume Sold	288 oz	23.53	63.60	0.01	2652
A	National advertising per brand	Mill of $	3.54	6.30	0.00	40.37
Distance	Distance to nearest brewing facility	Miles/1000	0.50	0.73	0.00	5.61

Source: IRI database, University of Connecticut; Arbitron-Leading National Advertisers.

Data Description

Table 1 provides a description and summary statistics of the variables used.[19] The main source is the Information Resources Inc. (IRI) Infoscan Database. The IRI data includes prices and total sales for several hundred brands for up to 58 cities over 20 quarters (1988-1992).[20] Volume sales (Quantity) in each city are reported as the number of 288-ounce units sold each quarter by all supermarkets in that city and price is an average price for a volume of 288 oz. for each brand. To maintain focus on brands with significant market share, all brands with a local market share of less than 3% are excluded from the sample. This selection criterion provides a sample of 64 brands produced by 13 different brewers. Appendix A contains a table of the brands chosen and their corresponding brewers. Advertising data (*A*) was obtained from the Leading National Advertising annual publication; these are quarterly data by brand comprising total national advertising expenditures for 10 media types.

[19] To conserve space, we have excluded variables used in the demand estimation. Details of these variables and a further description of the database and the data selection procedure can be found in Rojas (2006).

[20] The actual market definitions of these cities are broader than a single city and are usually referred to as "metropolitan areas". The term city here is used for simplicity. In general, definitions of these metropolitan areas are broader than those given by the BLS.

Table 2: Geographic Location of Brewing Plants by Brewer

Brewer	Cities where Brewing Plants are located (state/country)
Anheuser-Busch	Baldwinsville (NY); Columbus (OH); Fairfield (CA); Houston (TX); Jacksonville (FL); Los Angeles (CA); Merrimack (NH); Newark (NJ); St. Louis (MO); Tampa (FL); Williamsburg (VA)
Adolph Coors	Golden (CO); Elkton (VA); Memphis (TN)*
Bond/Heileman	Baltimore (MD); Frankenmuth (MI); LaCrosse (WI); Perry (GA); Phoenix (AZ); San Antonio (TX); St. Paul (MN);
Genessee	Rochester, NY
Grupo Modelo	Mexico D.F. (Mexico)
Goya	New York State (exact location unknown)
Heineken	Netherlands
Labatt	Edmonton, Winnipeg, Saint John (New Brunswick), St John's (Newfoundland); Etobicoke (Ontario); London (Ontario); Toronto (Ontario); Waterloo (Ontario); Saskatoon (Saskatchewan)
Molson	Calgary, Edmonton, Vancouver (BC), St John's (Newfoundland); Barrie (Ontario); Toronto (Ontario); Regina (Saskatchewan)
Pabst	Milwaukee (WI); Tampa (FL); Tumwater (WA)
Philip-Morris/Miller	Albany (GA); Eden (NC); Fulton (NY); Ft. Worth (TX); Irwindale, (CA); Milwaukee (WI)
Stroh	Longview (TX); Memphis (TN); St. Paul (MN); Van Nuys (CA); Winston-Salem (NC)
FX Matts	Utica (NY)

* Established in 1990

Location of Brewing Facilities and Distance

Shipping costs are proxied by the distance that a product has to travel from the brewery to a given city.[21] Brewers in our sample typically operate more than one brewing facility (table 2). The shipping distances (and thus shipping costs) in our data vary widely across brewers and cities and are large in many cases: nearly all brewers sell their brands in more than one city while most brewers sell their brands nationally.[22] The Distance variable in table 1 was computed as follows: for each brand in a city-quarter pair we calculated the distances to all brewing facilities

[21] Section 5.4 discusses the plausibility of distance as an appropriate proxy for shipping costs.
[22] Anheuser-Busch, Coors, Corona, Heineken, Labatt, Molson, Miller, Pabst, and Stroh sell brands that are present in a large portion of all cities (typically above 70% of all cities, with many brands present in above 90% of all cities). While few brewers sell their brands only in a particular region, they typically sell in more than 3 or 4 cities, which, given the city definitions in our data, represents a large geographic area.

of its corresponding brewer and chose the shortest distance.[23] Because advertising is only observed at the national level, we can not exploit shipping distances (only the change in the federal excise tax) in our analysis of advertising.

It is important to point out the economic importance of shipping costs in this industry. In our dataset, the average distance that a brand has to travel to a city is 508 miles. At an estimated \$2.68 per mile rate on an 80,000 lbs capacity truck (USDA, 2006), the average shipping cost is \$0.25/case, approximately 40% the magnitude of the federal excise tax increase of 1991.

4. Empirical Strategy

We employ a two-stage approach. In the first stage we adopt a structural methodology and estimate demand for the differentiated products to recover an estimate of brand equity/quality ($\hat{\xi}$). In the second stage we use a reduced from approach by regressing each of our three variables of interests (price, quantity and advertising) on several control variables and our measure of brand quality (estimated in the first stage) to determine if brands with different quality levels are affected differently by the federal tax increase of 1991 or by larger shipping costs. In this section we detail each step used in our strategy.

4.1. Demand Estimation

Demand estimation follows Rojas and Peterson (2008); here we present a shorter description of the method. Given data on observed quantity (q_{jt}) and price (p_{jt}) for brand j in market t (a city-quarter pair in this study), the linear approximation to the Almost Ideal Demand System (AIDS) of Deaton and Muelbauer is employed:

(4) $w_{jt} = a^*_{jt} + \sum_k b_{jk} \log p_{kt} + d_j \log(x_t / P_t)$

[23] Distance between two locations corresponds to the geographical distance (straight line) rather than the driving distance. This reduced the computational time of calculating the numerous plant-city distances.

$x_t = \sum_j p_{jt} q_{jt}$ are total expenditures, $w_{jt} = \left(p_{jt} q_{jt}\right) / x_t$ is brand $j's$ sales share and $\log P_t$ is a price index approximated the loglinear analogue of the Laspeyeres index: $\log P_t \approx \sum_j w_j^o \log(p_{jt})$; where w_j^o is brand $j's$ 'base' share, defined as $w_j^o \equiv M^{-1} \sum_{t=1}^{M} w_{jt}$.

Advertising for brand k (A_k) is incorporated into equation (4) by defining the intercept term as: $a_{jt}^* = a_{jt} + \sum_k c_{jk} A_{kt}^{\gamma}$. The parameter γ is included to account for decreasing returns to advertising. Substituting the redefined intercept into equation (4) and including an econometric error term gives:

(5) $w_{jt} = a_{jt} + \sum_k c_{jk} A_{kt}^{\gamma} + \sum_k b_{jk} \log p_{kt} + d_j \log(x_t / P_t) + v_{jt}$

Equation (5) is as a first-order approximation in prices and advertising to a demand function that allows unrestricted price and advertising parameters. In order to reduce the number of cross-price and cross-advertising coefficients that need to be estimated, we employ the Distance Metric method of Pinkse, Slade and Brett (2002). This method specifies each cross-coefficient (b_{jk} and c_{jk}) as a function of the distance between brands j and k in product space.

Distance measures may be either continuous or discrete. For example, alcohol content can be used to construct a continuous distance measure. Dichotomous variables that group brands into different market segments are used to construct discrete distance measures and take a value of 1 if brands j and k belong to the same grouping and zero otherwise. Continuous distance measures use an inverse measure of distance (closeness) between brands.

The terms b_{jk} and c_{jk} are specified as a linear combination of distance measures: $b_{jk} = \sum_{r=1}^{R} \lambda_r \delta_{jk}^r$, $c_{jk} = \sum_{s=1}^{S} \tau_s \mu_{jk}^s$, where $\delta_{jk} = \{\delta_{jk}^1, ..., \delta_{jk}^R\}$ is the set of distance measures for cross-prices and $\mu_{jk} = \{\mu_{jk}^1, ..., \mu_{jk}^S\}$ the set of measures for cross-advertising; λ and τ are the coefficients to be estimated. Replacing the definitions of b_{jk} and c_{jk} into (5) and regrouping terms gives the empirical demand equation:

$$(6)\ w_{jt} = a_{jt} + b_{jj}\log p_{jt} + c_{jj}A_{jt}^{\gamma} + \sum\nolimits_{r=1}^{R}(\lambda_r \sum\nolimits_k \delta_{jk}^r \log p_{kt}) + \sum\nolimits_{s=1}^{S}(\tau_s \sum\nolimits_k \mu_{jk}^s A_k^{\gamma}) + d_j \log(x_t / P_t) + v_{jt}$$

The reduced number of estimated coefficients λ_r and τ_s and the distance measures between brands (δ_{jk} and μ_{jk}) are used to obtain cross-terms (b_{jk} and c_{jk}). To reduce the number of parameters to be estimated, symmetry of the Slutsky matrix (i.e. $b_{jk} = b_{kj}$ and $c_{jk} = c_{kj}$) is imposed by setting λ and τ to be equal across equations. In addition, it is assumed that the own-price and own-advertising coefficients (b_{jj} and c_{jj}), and the price index coefficient (d_j), are equal across equations thereby reducing estimation to one equation (since symmetry is imposed). To relax this last assumption, the coefficients b_{jj}, c_{jj}, and d_j are specified as linear functions of brand $j's$ characteristics.

Brand equity

As pointed out by Berry (1994), the error term v_{jt} in (5) may contain a product-specific variable (ξ_{jt}) that can be interpreted as the mean of consumers' valuations of an unobserved product characteristic such as product quality (or brand equity as interpreted here).[24] Of importance in Berry's analysis is the potential source of endogeneity of ξ_{jt} since it can be correlated with prices. To account for this source of endogeneity one can use instrumental variables. In addition, one can write v_{jt} as $v_{jt} = \xi_{jt} + e_{jt}$, where e_{jt} is a mean-zero stochastic term, and further decompose ξ_{jt} into a market-invariant part ξ_j and its market-specific deviation

[24] Berry's discussion is in the context of the random utility model. However, there is no apparent reason for this analysis not to be valid under the assumed functional form.

$\Delta\xi_{jt}$ (i.e. $v_{jt} = \xi_j + \Delta\xi_{jt} + e_{jt}$). Then, the inclusion of brand fixed effects (in addition to instrumentation) in the estimation of (5) not only reduces the endogeneity bias caused by ξ_{jt}, but it also improves the fit of the model. A by-product that is of key relevance to our analysis is that the estimates on the brand fixed effects $\hat{\xi}_j$ can be used as a measure of brand equity.[25]

By using brand equity as a proxy for quality we avoid using subjective measures of quality. Instead, the data informs us about a brand's valuation regardless of how this valuation has been created (deceptive advertising, word of mouth, etc.). Importantly, as with quality, brand equity is an unambiguous measure: it can generally be considered as a dimension of the product through which brands are vertically differentiated as required in the Alchian-Allen and Barzel analyses.

Endogeneity

Our estimation of (5) deals with the endogeneity of prices and advertising in several ways. Two different identification assumptions are utilized: after controlling for brand, city, and time specific effects, demand shocks are independent across cities and across time. We employ both assumptions to address price endogeneity and the second to address the endogeneity of advertising. Independence of demand shocks across cities allows us to use prices of a brand in other cities as an instrument, whereas independence of demand shocks across time allows us to use lagged prices (and lagged advertising) as instruments.[26] A full discussion of our treatment of endogeneity can be found in Rojas and Peterson; here we outline the general approach.

[25] We should interpret the estimated values with care since they are not directly measured in monetary terms. Here, the estimate of ξ_j measures the additional sales shares (w_j) that a brand's equity is responsible for. This is a "rescaled" version of brand equity as a shifter of volume sales (q_j).

[26] Because brand-level advertising expenditures are invariant across markets, we can not utilize the first identifying assumption to construct an additional instrument for advertising.

The main idea is to conduct robustness checks of demand estimates by comparing estimates from the two different identifying assumptions. In addition, we conduct a test of instrument exogeneity that is similar to a GMM differential test in which the error-orthogonality of a subset of "suspicious" instruments is tested. We conduct two tests. In one test the suspicious instruments are those constructed with lagged prices and another in which suspicious instruments are those constructed with prices in other cities.

We adopt other measures to improve the validity of our identifying assumptions and to reduce other endogeneity sources. National advertising expenditures and time dummy variables reduce the potential that demand shocks may be correlated across markets, whereas the inclusion of brand dummies control for market-invariant unobserved product characteristics that can be correlated with prices (as described earlier). In addition, we create proxies for city-specific marginal costs that are used as additional price instruments to conduct overidentifying tests. Finally, since expenditures, (x_t), are constructed with price and quantity variables, this term is instrumented with median income.

One last feature of our data increases our confidence in one of the identifying assumptions. The IRI data is based on broadly defined city/regional markets which further reduce the possibility of potential correlation between the unobserved shocks across markets.

4.2. Estimation of Quantity, Price and Advertising Effects

For empirical purposes, we rewrite equation (1) in elasticity format:

(1') $\frac{z_j}{z_k} T\left(\varepsilon_{z_j,T} - \varepsilon_{z_k,T}\right) > 0$

where $\varepsilon_{z,T} = \dfrac{\partial z}{\partial T}\dfrac{T}{z}$. Thus, a sufficient condition for the existence of Alchian-Allen and Barzel effects is that $\left(\varepsilon_{z_j,T} - \varepsilon_{z_k,T}\right) > 0$. A simple log-log OLS specification can be used to directly test this condition:

(6) $\log z_{jt} = \alpha_{jt} + \theta \cdot \xi_j + \rho_1 \cdot \log T^{\$} + \rho_2 \cdot \log T^{\$} \cdot \xi_j + \gamma_1 \cdot \log T^{miles} + \gamma_2 \cdot \log T^{miles} \cdot \xi_j$

where α_{jt} contains a variety of fixed effects, $T^{\$}$ is equal to 65 cents for the pre-tax-increase period (1988-1990) and equal to 130 cents for the post-tax-increase period (1991-1992), and T^{miles} is the shipping distance in miles.[27] The coefficients of interest are ρ_2 and γ_2 because they measure (after all other factors have been accounted for) whether quantity, price and advertising of brands with different qualities respond differently to a change in a per-unit charge. The elasticity $\varepsilon_{z_j,T}$ is given by: $\rho_1 + \rho_2\xi_j + \gamma_1 + \gamma_2\xi_j$, hence positive estimates of ρ_2 and γ_2 suggest that the condition $\left(\varepsilon_{z_j,T} - \varepsilon_{z_k,T}\right) > 0$ holds for $\xi_j > \xi_k$. However, since quality is not a dichotomous variable, statistically significance of the coefficients ρ_2 and γ_2 alone does not guarantee a statistically significant difference of $\left(\varepsilon_{z_j,T} - \varepsilon_{z_k,T}\right)$. Instead, for $\xi_j > \xi_k$ to hold, the following condition(s) should be satisfied for a one-tailed test at the 95% confidence level:[28]

(7) $\hat{\rho}_2 \times \hat{\xi}_j - 1.96 \times (SE_{\hat{\rho}_2} \times |\hat{\xi}_j|) \geq \hat{\rho}_2 \times \hat{\xi}_k + 1.96 \times (SE_{\hat{\rho}_2} \times |\hat{\xi}_k|)$

(8) $\hat{\gamma}_2 \times \hat{\xi}_j - 1.96 \times (SE_{\hat{\gamma}_2} \times |\hat{\xi}_j|) \geq \hat{\gamma}_2 \times \hat{\xi}_k + 1.96 \times (SE_{\hat{\gamma}_2} \times |\hat{\xi}_k|)$

[27] We separate the effect of shipping costs and the excise tax to study potentially different effects from these two types of per-unit charges.

[28] Given the large number of observations, we assume that the distribution of the estimated coefficients is well approximated by a normal distribution. The absolute value of the quality measure (in parenthesis) is employed because there are some values of $\hat{\xi}$ that are negative.

where SE denotes the standard error of the estimated coefficient. We check whether this condition holds for each of the $J(J-1)/2$ =2,016 possible pair-wise comparisons of quality levels.

In principle, one could also use state excise taxes to analyze the problem at hand. In addition, for comparison purposes, one could study whether the effect of state ad-valorem taxes on sales is different than that of excise taxes, as predicted by Barzel, for example. However, the use of state taxes (both ad-valorem and excise) is problematic because their exogeneity is suspect: state-level idiosyncrasies are likely to be correlated with state taxes as well as with sales (price and quantity). We collected data on these two types of taxes at the state level and the results obtained appeared difficult to interpret. Specifically, results tended to be inconsistent with the results we obtained using federal tax and shipping costs; we attributed this problem to the potential endogeneity of state-level taxes.

4.3 Correcting for Heteroskedastic and Correlated Errors

Our data structure has the potential of having errors that are correlated across time and space. Spatial correlation has two dimensions: across cities and across brands. We compute robust standard errors by implementing a covariance matrix estimator that accommodates for heteroskedasticity and for both types of correlation. In the context of GMM estimation (used for our demand equation), this is done by properly assigning non-zero entries to the GMM weighing matrix: to the diagonal elements (to correct for heteroskedasticity) and to the off-diagonal elements that are suspected to be correlated across time and space (Newey and West, 1987; Conley, 1999). The non-zero entries correspond to the product of residuals from the first stage in a two-stage least squares procedure. In OLS estimation (our second step), only the standard errors are modified via a sandwich estimator for the variance-covariance matrix where the non-zero elements of the weighing matrix are defined in a similar way as the GMM weighing matrix, except that the residuals come from OLS estimates.

Specifically, we assign non-zero off-diagonal entries to correct for temporal correlation of one period and for brands that are nearest neighbors in product space (see Rojas, 2005, for details on product space definition and computation of nearest neighbor measures). Also, non-zero entries are assumed across cities that are located in the same region of the US according to the Bureau of Labor Statistics classification (West, Midwest, South and Northeast).[29] We developed our own computer program to carry out this procedure as there are no statistical packages with this feature for a three-dimensional panel.

The importance of this correction can not be easily overstated. In our application below, all standard errors of the estimated coefficients increase (frequently in the range of 80% to 100%) when they are corrected for heteroskedasticity and autocorrelation, with the latter having a larger impact. After this robustness correction, several coefficients ceased to be statistically significant. These results are consistent with the large standard error biases that arise when spatial dependence is not accounted for (Driscoll and Kraay, 1998).

5. Results

5.1 Quality Estimates

Table 3 reports the estimates of brand equity that are recovered from demand estimation, ordered from highest to lowest. The obtained values are with respect to the omitted brand in the demand estimation (in this case Utica), which has an equity value of zero.[30] The values are reasonable.

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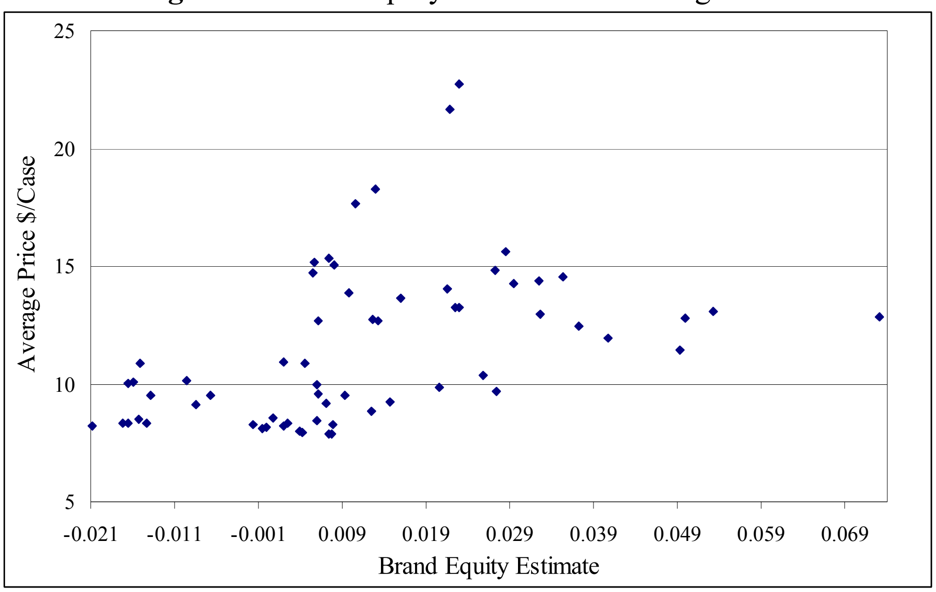
[29] Tests on estimates of the residuals from the demand equation and from the reduced from equations without this correction suggested the presence of temporal and spatial error correlation. There are multiple spatial distance definitions that can be used when assigning non-zero entries to the covariance matrix (see Conley, 1999); we use a specific discrete distance measure, but our main results are robust to other distance definitions (e.g. Euclidean distance, nearest neighbor, common boundary neighbor).

[30] Changing the excluded brand does not change our results.

For example, the four best selling brands, Budweiser, Bud Light, Coors Light and Miller Lite (with market shares of 19%, 6%, 7% and 9%, respectively) have the four largest estimates of brand equity. In a market with many dozens of brands, it is remarkable for four brands to account for 41% of sales; our estimates of brand equity are consistent with (and may explain) this observation. Imported brands, which would be expected by many consumers to be of superior quality, also rank relatively high on the list: Labatt's Blue is 10th, Heineken is 16th and Corona is 19th. This is also true for beers classified in the "super-premium" segment such as Henry Weinhard Ale or Michelob Golden Draft. Conversely, budget beers, such as Schaefer or Milwaukee's Best, are relegated to the end of the list.

Figure 2: Brand Equity Estimate vs. Average Price



Note: Budweiser is not included.

Higher quality is typically associated with a higher price; thus, checking if our measure of quality is positively related with price can help us better assess the plausibility of our proxy for quality. Figure 2 shows depicts a positive relationship between price and the brand equity estimate (we exclude the Budweiser outlier in the graph to better visualize the relationship). A statistically significant correlation coefficient of 0.37 confirms the graphical pattern. We acknowledge the fact that brand equity may also proxy for other product attributes that are difficult to measure (e.g. "status"), but the plausibility of our estimates provides us with some confidence that they are reasonably correlated with brand quality.

Table 3: Estimates of Brand Equity ($\hat{\xi}$) Obtained From Demand Equation

Brand (Brewer)	**Quality Estimate**	**Brand**	**Quality Estimate**
Budweiser (AB)	0.1571	White Stag (B)	0.0077
Miller Lite (PM)	0.0731	Michelob Dry (AB)	0.0074
Budweiser Light (AB)	0.0532	Goebel (S)	0.0074
Coors Light (AC)	0.0499	Schmidts (B)	0.0070
Genesee (GE)	0.0493	Coors Extra Gold (AC)	0.0061
Goya (GO)	0.0407	Natural Light (AB)	0.0061
Rainier (B)	0.0373	Piels (S)	0.0059
Henry Weinhard Ale (B)	0.0354	Matts (W)	0.0059
Old Vienna (M)	0.0326	Michelob Light (AB)	0.0057
Labatt's Blue (LB)	0.0325	Rolling Rock (LB)	0.0056
Michelob Golden Draft (AB)	0.0295	Busch Light (AB)	0.0046
Molson (M)	0.0285	Heidelberg (B)	0.0042
Old Style (B)	0.0274	Blatz (B)	0.0039
Henry Weinhard Private Res (B)	0.0273	Red White & Blue (P)	0.0025
Busch (AB)	0.0258	Meister Brau Light (PM)	0.0021
Heineken (H)	0.0230	Kingsbury (B)	0.0020
Genuine Draft (PM)	0.0229	Weidemann (B)	0.0008
Lone Star Light (B)	0.0224	Utica Club (W)	0.0000
Corona (GM)	0.0218	Falstaff (P)	-0.0006
Odoul's (AB)	0.0214	Milwaukee's Best (PM)	-0.0017
Lone Star (B)	0.0205	Old Milwaukee Light (S)	-0.0068
Budweiser Dry (AB)	0.0159	Old Milwaukee (S)	-0.0085
Old Style Light (B)	0.0148	Keystone Light (AC)	-0.0095
Miller High Life (PM)	0.0132	Pabst Blue Ribbon (P)	-0.0139
Molson Golden (M)	0.0130	Hamms (P)	-0.0143
Coors (AC)	0.0126	Strohs (S)	-0.0151
Kochs (GE)	0.0124	Black Label (B)	-0.0153
Labatt (LB)	0.0106	Schlitz (S)	-0.0159
Miller Genuine Draft Light (PM)	0.0097	Keystone (AC)	-0.0165
Sterling (B)	0.0092	Meister Brau (PM)	-0.0166
Michelob (AB)	0.0080	Olympia (P)	-0.0172
Hamms Light (P)	0.0078	Schaefer (S)	-0.0209

AB=Anheuser-Busch; AC=Adolph Coors; B=Bond Corporation; GE=Genesee; GM=Grupo Modelo; GO=Goya; H=Heineken; LB=Labatt; M=Molson; P=Pabst; PM=Philip Morris/Miler; S=Stroh; W=FX Matts.

5.2 Alchian-Allen and Barzel Effects

The main results, using estimable equation (6), are reported in table 4. The table is split in three sections, each reporting regressions for each of the dependent variables of interest: quantity, price and advertising. For quantity and price regressions we include time, city and brewer fixed effects,[31] and consider two specifications one with and the other without the shipping cost proxy (distance). Because advertising expenditures are only available at the national level, including the distance variable is not relevant in the advertising regression; for the same reason, in this regression we include time and brewer fixed effects, but not city fixed effects.

Our estimate of quality ($\hat{\xi}$) is significant in all specifications and has the expected sign: higher quality brands have higher volume sales and prices, and are advertised more intensely. The effect of the federal excise tax increase is not statistically significant in either specification of the quantity regression, but it appears to have increased average prices as well as advertising expenditures. The interaction of the federal excise tax and quality is only statistically significant in the advertising regression; these two last results suggest that after the 100% tax increase, advertising intensity increased for all brands but by a larger amount for higher quality brands.

[31] We do not include brand fixed effects to avoid collinearity with the quality variable.

Table 4: OLS Regressions of the Effect of Brand Quality on Equilibrium Quantity, Advertising Expenditures and Prices as a Result of the Federal Tax Increase of 1991 and Shipping Costs (Distance)

	Dependent Variable				
Explanatory Variables [equation (6) notation]	Log (Quantity)		Log (Price)		Log (Advertising)
	Spec. 1	Spec. 2	Spec. 1	Spec. 2	Spec. 1
Quality [$\hat{\xi}$]	17.61* (0.52)	15.73* (0.97)	1.44* (0.13)	0.57* (0.13)	37.82* (0.79)
Log (Federal Excise Tax) [$\hat{\rho}_1$] = \$0.65 /288 oz 1988-90 = \$1.30 /288 oz 1991-92	0.20 (0.14)	0.19 (0.14)	0.19* (0.03)	0.19* (0.03)	0.48* (0.13)
Log (Federal Excise Tax) x **Quality [$\hat{\rho}_2$]**	**0.99** **(1.27)**	**1.05** **(1.25)**	**0.16** **(0.27)**	**0.15** **(0.27)**	**11.30*** **(1.79)**
Log (Distance to Brewer) [$\hat{\gamma}_1$]		-0.11* (0.01)		-0.004* (0.002)	
Log (Distance to Brewer) x Quality [$\hat{\gamma}_2$]		**0.38**** **(0.17)**		**0.18*** **(0.03)**	
R-squared	*0.49*	*0.50*	*0.69*	*0.69*	*0.55*

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

Notes: i) Number of observations: 33,892.

ii) In parenthesis are standard errors, corrected for heteroskedasticity as well as temporal and spatial autocorrelation (see section 4.3 for details of this computation).

iii) Quantity and Price regressions include firm, city and time fixed effects. Advertising regression includes firm and time fixed effects.

The variable Distance is significant in both quantity and price regressions, but has the expected sign only in the quantity regression. Note, however, that the unexpected coefficient for Distance in the price regression is economically small. The interaction of Distance with the quality variable, on the other hand, has the expected sign and is statistically significant in both quantity and price regressions. This latter result suggests that higher per-unit shipping costs increase quantity and price by a larger amount for a higher quality brand than for a lower quality brand. Finally, all regressions fit the data reasonably well: the R-squared measure is close or above 0.50.

As indicated in section (4.2), statistical significance of the coefficients for the interactions of interest is not enough to guarantee the existence of the effects of interest: $\left(\varepsilon_{z_j,T} - \varepsilon_{z_k,T}\right) > 0$. Using the estimated coefficients and corresponding standard errors of the interaction terms in table 4, together with our quality estimates, we test the statistical significance of $\left(\varepsilon_{z_j,T} - \varepsilon_{z_k,T}\right)$ for all 2,016 possible pairs of brands; we employ condition (8) for quantity and price regressions and condition (7) for the advertising regression. The quantity effect $\left(\varepsilon_{q_j,T} - \varepsilon_{q_k,T}\right) > 0$ is statistically significant for 800 of the pair-wise comparisons (40% of the time). Price effects $\left(\varepsilon_{p_j,T} - \varepsilon_{p_k,T}\right) > 0$ and advertising effects $\left(\varepsilon_{A_j,T} - \varepsilon_{A_k,T}\right) > 0$ are much more widespread (consistent with the high statistical significance of the corresponding interactions): 1,581 (78%) and 1,719 (85%) of all cases, respectively.

Of all three effects, the advertising effect, $\left(\varepsilon_{A_j,T} - \varepsilon_{A_k,T}\right) > 0$, is the largest in economic terms: the coefficient on the interaction term (Log (Federal Excise Tax) x Quality [$\hat{\rho}_2$]) is two orders of magnitude larger than in the price and quantity regressions. Given the ambiguous effects predicted by our model, these results can be interpreted as generally supportive of the predictions by the Alchian-Allen and Barzel theories, with the price effect being smaller but more widespread than the quantity effect.

5.3 Plausibility of Assumptions and Robustness Checks

Distance as a Proxy for Shipping Costs

Since we employ distance as our proxy for shipping costs, we are assuming that all brewers incur in a fixed shipping cost per mile, regardless of distance or of shipment size. This may be a strong assumption since transportation companies may offer a discount for larger distances or for larger shipments.[34] The best data source available for distance discounts is an index constructed by the USDA for the years 2001-2007; this index consists of average shipping rates (for grain transportation in trucks) computed for three types of distances: 25 miles, 100 miles and 200 miles.[35] There is a large difference in shipping rates between 25 and 100 miles ($4.70 vs. $2.68 per mile-container) and a relatively smaller one between 100 and 200 miles ($2.68 vs. $2.14 per mile-container), which suggests that per-mile shipping rates flatten out beyond 200 miles. In our data, the shipping distance is greater than 100 miles for 88% of the observations and is greater than 200 miles for 75% of the observations, which suggests that our assumption of constant shipping cost per-mile regardless of distance may not be too strong. To be sure, table 5.1 shows the regression results that only included observations with shipping distances greater than 100 or 250 miles; as can be seen in the table, our qualitative findings remain unchanged and even appear to be of larger magnitude. Excluding imports (which are

[34] Since the shipping companies incur in some fixed costs, regardless of distance, the per-mile rate may be higher for shorter trips.
[35] We confirmed, via phone interview with the USDA, that data on this or other similar measures do not exist for the US prior to 2001. In 2001, the USDA implemented a quarterly nationwide survey that was specifically designed to construct this index, see http://www.ams.usda.gov/tmdtsb/grain/Archive.htm.

subject to different modes of transportation) did not change our main results either (table 5.1), except for the quantity effect which ceased being statistically significant.

Table 5.1: Robustness Checks for Table 4 estimates (Distance Assumption)

	[Robustness criteria] / Dependent Variable					
Explanatory Variables	[Distance > 100 miles]		[Distance > 250 miles]		[Domestic Beers Only]	
	Log (Q)	Log (P)	Log (Q)	Log (P)	Log (Q)	Log (P)
Quality [$\hat{\xi}$]	9.72*	-2.43*	0.37	-2.29	16.21*	0.54*
	(4.09)	(0.74)	(8.27)	(1.64)	(1.00)	(0.14)
Log (Federal Excise Tax)	0.18	0.19*	0.24	0.19*	0.43*	0.20*
	(0.14)	(0.03)	(0.15)	(0.02)	(0.15)	(0.03)
Log (Federal Excise Tax) x $\hat{\xi}$	**1.14**	**0.19**	**1.00**	**0.17**	**1.15**	**0.18**
	(1.34)	**(0.30)**	**(1.53)**	**(0.35)**	**(1.26)**	**(0.27)**
Log (Distance)	-0.23*	-0.01*	-0.22*	-0.01	-0.10*	-0.01*
	(0.04)	(0.005)	(0.06)	(0.008)	(0.01)	(0.002)
Log (Distance) x $\hat{\xi}$	**1.39****	**0.72***	**2.94****	**0.71***	**0.29**	**0.19***
	(0.72)	**(0.13)**	**(1.39)**	**(0.28)**	**(0.18)**	**(0.03)**
R-squared	*0.49*	*0.72*	*0.52*	*0.76*	*0.47*	*0.54*
# Obs.	29,309		21,784		30,187	

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

Notes: i) In parenthesis are standard errors, corrected for heteroskedasticity as well as temporal and spatial autocorrelation (see section 4.3 for details of this computation).

ii) All regressions include firm, city and time fixed effects.

To the best of our knowledge, there is no public data available on volume discounts for shipping rates. However, since all brands included in the data are shipped in large quantities (i.e. are not craft beers), we suspect that our assumption of a fixed shipping rate regardless of volume is not be too restrictive.

Fixed and Exogenous Quality Assumption

In our analysis we assume that brand quality remains constant over time, but it is possible that quality adjustments occur as a result of larger per-unit charges. To check the plausibility of our assumption, we estimated an alternative specification of demand that included "brand-shift" dummy variables, where "shift" refers to a dichotomous variable that takes a value of 1 for the post-tax increase period (1991-1992) and zero otherwise. This procedure generates two sets of brand equity estimates, one for the pre-tax increase period and another for the post-tax increase period. Using a Wilcoxon signed-ranks test for matched-pairs we failed to reject the hypothesis that brand equity estimates shifted upwards (as predicted by Barzel) as a result of the tax increase.[36] We also carried out quantity and price regressions with data from the post-tax increase period only as well as with data from the pre-tax increase period; the results are shown in table 5.2. Except for the quantity regression in the 1991-1992 period, our main results are confirmed.

Also, our proxy for quality is an estimate of brand equity, but an important determinant of brand equity is advertising. Thus, the validity of estimates in the advertising regression may be compromised if advertising expenditures significantly affected brand equity during the 1988-1992 period. Brand equity is typically built over many years, so our results are valid to the extent that our five year period is a "snapshot" of this long process; this argument is more plausible for those brands that have been in the market for many years, as equity growth stabilizes over time. On the other hand, this argument may be problematic for newer brands since they are more likely to be subject to significant brand equity "build up" early on, especially through advertising. In general, most of the brands in our sample have been in the market for decades (Appendix A), but

[36] A t-test produced a similar conclusion.

to be sure we checked the robustness of our results when newer brands are excluded from the regressions.

Table 5.2 reports regressions with two subsets of brands using different cutoff introduction dates: 1970 and 1980. The reason for these two cutoff points is that there are three important time periods of brand introductions in our data set. Most brands were introduced prior to 1970's (42 out of 64) with many of them being introduced as early as the 19th century. During the 1970's there is an important wave of brand introductions with the appearance of the light segment (9 out of 64). Finally, the introduction of several successful brands (and other light beers) occurred after 1980 (13 out of 64). The results in table 5.2 indicate that the advertising effect is still present, but it is approximately 32%-35% smaller. The quantity effect, however, is no longer significant and the price effect is only significant in one of the specifications (and smaller in magnitude).

We would like to discuss a potential limitation of the robustness checks presented in this section. While these specifications focus on a subset of the data that is less likely to be sensitive to our distance and quality assumptions, they also exclude important information that may be critical in assessing the effects of interest. For example, by excluding imports, we are excluding a number of beers with relatively high quality levels thereby reducing our ability to capture the effect of consumers' substitution towards those brands when per-unit charges are larger. In our case, however, this potential limitation turns out to be good news: our main results are *still* present in subsamples that restrict the product set actually available to consumers.

Table 5.2: Robustness Checks for Table 4 estimates (Quality Assumption)

Explanatory Variables	[1988-1990] Log (Q)	[1988-1990] Log (P)	[1991-1992] Log (Q)	[1991-1992] Log (P)	[Introduction Date < 1970] Log (Q)	[Introduction Date < 1970] Log (P)	[Introduction Date < 1970] Log (A)	[Introduction Date < 1980] Log (Q)	[Introduction Date < 1980] Log (P)	[Introduction Date < 1980] Log (A)
$\hat{\xi}$	14.77* (1.15)	0.53* (0.17)	16.86* (1.56)	0.55* (0.19)	11.84* (1.05)	0.46* (0.16)	25.76* (0.63)	14.26* (0.98)	0.80* (0.14)	38.12* (0.75)
Log (Federal Tax)					0.18 (0.17)	0.20* (0.03)	0.31** (0.16)	0.45* (0.15)	0.20* (0.03)	1.23* (0.16)
Log (Federal Tax) x $\hat{\xi}$					**0.48** **(1.20)**	**-0.04** **(0.16)**	**7.25*** **(1.23)**	**0.42** **(1.17)**	**0.01** **(0.23)**	**7.67*** **(1.71)**
Log (Distance)	-0.11* (0.01)	-0.004** (0.002)	-0.10* (0.02)	-0.004** (0.002)	-0.13* (0.02)	0.000 (0.002)		-0.11* (0.01)	-0.002 (0.002)	
Log (Distance) x $\hat{\xi}$	**0.50*** **(0.21)**	**0.17*** **(0.03)**	**0.18** **(0.28)**	**0.19*** **(0.04)**	**0.15** **(0.18)**	**0.03** **(0.02)**		**0.09** **(0.18)**	**0.10*** **(0.02)**	
R-squared	*0.51*	*0.69*	*0.50*	*0.68*	*0.49*	*0.77*	*0.51*	*0.51*	*0.72*	*0.54*
# Obs.	18,369		15,523		20,750			26,009		

[Robustness criteria] / Dependent Variable

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

Notes: i) In parenthesis are standard errors, corrected for heteroskedasticity as well as temporal and spatial autocorrelation (see section 4.3 for details of this computation).

ii) Quantity and Price regressions include firm, city and time fixed effects. Advertising regression includes firm and time fixed effects.

5.4 Implications for Public Policy

Policy makers may care to know about the consequences for alcohol consumption and tax revenues when brands of different qualities respond differently to a per-unit tax. Our results indicate that the federal tax may not have an effect on quantity (in table 4, $\hat{\rho}_1$ is not statistically different from zero). Furthermore, the federal tax increase does not have a differential effect on the sales of brands with different quality levels ($\hat{\rho}_2$ is not statistically different from zero). However, we do find that another type of per-unit charge (shipping costs) possesses these effects ($\hat{\gamma}_1$ and $\hat{\gamma}_2$ are statistically significant). We hypothesize that the richer (cross-sectional) variation in the distance variable allows for a better identification of the effects of interest. Therefore, for the purposes of this section, we assume that the effect of a larger per-unit tax on quantity is captured by coefficients $\hat{\gamma}_1$ and $\hat{\gamma}_2$. These coefficients can give us a ballpark estimate of the effects of quality heterogeneity on beer consumption and, thus, on tax revenue.[37]

A simple estimate of the percentage change in quantity consumed as a result of a given percentage change in a per-unit tax can be approximated with the simple formula:

$$\tilde{\varepsilon}_{q,T} = \hat{\gamma}_1 + \sum_{j=1}^{J} \varpi_{jy} \cdot \hat{\xi}_j \cdot \hat{\gamma}_2$$

where ϖ_j is a weight given by the market share of brand *j* in year *y*. In the absence of quality effects ($\hat{\gamma}_2 = 0$), a 10% increase in the per-unit tax on beer reduces consumption by 1.1% ($\hat{\gamma}_1 = -0.11$). When quality matters ($\hat{\gamma}_2 = 0.38$), the elasticity estimate ($\tilde{\varepsilon}_{q,T}$) is

[37] The inference in this section should, of course, be interpreted with caution as the quantity effect in our regressions is the least significant (and less robust) of all three effects.

$\tilde{\varepsilon}_{q,T} = -0.11 + 0.02 = -0.09$.[38] Thus, in the presence of quality heterogeneity, the effect of a per-unit tax on beer consumption is smaller by approximately two percentage points, or $18\% = \frac{2\%}{11\%}$.

This, in turn, means that tax revenues would be higher in comparison to the case when products are of homogeneous quality. With total US consumption of beer currently estimated at 210 million barrels per year (Beer Institute, 2007), the "quality premium", $\sum_{j=1}^{J} \varpi_{jy} \cdot \xi_j \cdot \hat{\gamma}_2$, given a 100% increase in the federal excise tax would account for an additional consumption of 4.2 million barrels per year, or $75.6 million of additional tax revenues (based on the current $18/barrel federal excise tax).

6. Conclusion

Our objective in this paper was to empirically explore how tax incidence analysis is affected by the presence of vertically differentiated products in the context of per-unit taxes. Specifically, we test the predictions implied by two popular models that study the effects of per-unit charges. First, we show that when the original models are modified to include more realistic market features, the predictions of these models become ambiguous. Surprisingly, however, we find important empirical support for the theoretical predictions of the two models.

Our evidence suggests that when per-unit charges are larger, equilibrium price and quantity increase relatively more for higher quality products than for lower quality products. These results are good news for tax collectors, but they also imply that a reduction in alcohol consumption will be less effective than in the absence of quality effects. Our estimates, together with prior empirical work on societal costs of alcohol consumption (see for example Chesson,

[38] The estimate of $\sum_{j=1}^{J} \varpi_{jy} \cdot \xi_j \cdot \hat{\gamma}_2$ corresponds to the year of 1992 (0.0187 to be precise). Estimates for other years are very similar, ranging from 0.0170 to 0.0187.

Harrison and Kassler, 2000; Markowitz and Grossman, 1998; and Ruhm, 1996), can prove to be useful to policy makers when assessing the trade-offs of tax policy in the beer market.

The quantity and price effects imply that consumer preferences shift towards higher quality products, and that quality can offset the negative effect of the per-unit tax on sales. This is good news for firms selling high quality brands: the negative impact of a per-unit tax on profits decreases as product quality increases. But the "tax shelter" that quality provides is offset by larger advertising expenditures: the per-unit tax increase causes advertising expenditures to increase more heavily for high quality than for low quality products. We note, however, that the "quality effect" is more important (i.e. more robust to alternative specifications and more significant) in determining price and advertising levels than in determining quantity.

We attempted to explore other interesting hypotheses. For example, a higher per-unit tax may shift consumption and production towards beers of higher alcohol content, but our results were not conclusive. Part of the reason for this result may be that, unlike cigarettes (see Evans and Farrelly, 1998), beer is not the only source of alcohol and consumers are able shift consumption towards other liquor products.[39] Another reason may be that alcohol content across beers tends to be rather homogeneous.

Our work is not without limitations. The plausibility of the results relies partly on quality being constant and also on our assumption of firms' choosing prices and advertising levels given consumers' perception of quality. There is some confidence that quality estimates did not shift upwards as a result of the 1991 tax increase, but this may not be true beyond 1992 (the last year of data availability). In the years following the tax increase of 1991, and most notably in the last decade, there has been an important surge in craft and import beers, which are deemed by many

[39] There is also some evidence about the substitutability between alcohol and illegal substances such as marihuana (e.g. DiNardo and Lemieux, 2001).

to be of superior quality. Unfortunately, our data does not capture this important trend, but we suspect that the recent popularity of "high quality" beers may have been fueled, at least in part, by the 1991 increase in the federal excise tax.

Our estimate of the effect of quality heterogeneity on beer consumption should be taken with some caution as well, as it is relies on the ceteris paribus assumption that equilibrium quantities, prices and advertising would be the same if quality heterogeneity were absent. Despite our limitations, and given the ambiguity of Alchian-Allen and Barzel effects in complex market environments, the results presented here provide relatively predictable evidence of how quality matters and that brand heterogeneity may have more profound implications when markets have other societal ramifications.

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Appendix A: Selected Brands by Brewer (Acronym and Country of Origin), [year of introduction]

Brewer	Brand	Brewer	Brand
Anheuser-Busch:	Budweiser [1876]	Grupo Modelo:	Corona [1926]
(AB, U.S.)	Bud Dry [1989]	(GM, Mexico)	
	Bud Light [1982]	Goya (GO, U.S.):	Goya [1936] [b]
	Busch [1955]	Heineken:	Heineken [1863]
	Busch Light [1989]	(H, Netherlands)	
	Michelob [1896]	Labatt:	Labatt [1992]
	Michelob Dry [1988]	(LB, Canada)	Labatt Blue [1951]
	Michelob Golden Draft [1991]		Rolling Rock [1939]
	Michelob Light [1978]	Molson:	Molson [1786]
	Natural Light [1977]	(M, Canada)	Molson Golden [1954]
	Odoul's [1990]		Old Vienna [1947] [b]
Adolph Coors:	Coors [1874]	Pabst:	Falstaff [1899]
(AC, U.S.)	Coors Extra Gold [1980]	(P, U.S.)	Hamms [1865]
	Coors Light [1978]		Hamms Light [1978] [b]
	Keystone [1989]		Olympia [1896]
	Keystone Light [1989]		Pabst Blue Ribbon [1895]
Bond Corp[a]:	Black Label [1840]		Red White & Blue [1920] [b]
(B, U.S.)	Blatz [1851]	Miller/Phillip Morris:	Genuine Draft [1985]
	Heidelberg [1900]	(PM, U.S.)	Meister Brau [1933]
	Henry Weinhard Ale [1856][b]		Meister Brau Light [1984]
	Henry Weinhard P. R. [1976]		MGD Light [1991]
	Kingsbury [1934]		Miller High Life [1903]
	Lone Star [1940]		Miller Lite [1975]
	Lone Star Light [1978]		Milwaukee's Best [1895]
	Old Style [1902]	Stroh:	Goebel [1873]
	Old Style Light [1978] [b]	(S, U.S.)	Old Milwaukee [1849]
	Rainier [1878]		Old Milwaukee Light [1978] [b]
	Schmidts [1844]		Piels [1883]
	Sterling [1913] [b]		Schaefer [1842]
	Weidemann [1870] [b]		Schlitz [1849]
	White Stag [1851]		Stroh [1850]
Genesee:	Genesee [1878]	FX Matts:	Matts [1950]
(GE, U.S.)	Kochs [1934]	(W, U.S.)	Utica Club [1934] [b]

[a] These brands correspond to G. Heileman Brewing Co., which was acquired in 1987 by Australian Bond Corporation Holdings; it is classified as a domestic brewer because this foreign ownership was temporary.
[b] Precise date not available; year approximated by authors using several industry sources.