Private Politics and Market Mediation

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Abstract

Private politics, or consumer activism, has proven effective in changing the behavior of private economic agents. In many markets, however, intermediaries stand between the consumer and the target of the activism, potentially reducing the effectiveness of private political action. We examine this effect for a recent consumer boycott of microbeads in the US toothpaste market. Relative to non-beaded toothpaste products, we demonstrate that consumer demand for microbeaded toothpaste decreased during the boycott period, but became more elastic, prompting retailers to offer selective price discounts on microbeaded toothpaste products. As a result, the equilibrium quantity of microbeaded toothpaste sold by retailers increased in the period immediately following the boycott.

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1 Introduction

Private politics applied through consumer activism is a pervasive phenomenon in industrialized countries. Consumer activists often rely on the use or threat of boycotts to place political pressure on private economic agents, a practice that has led timber producers to phase out production from old growth forests, fisheries to impose sustainable limits on catch, and fast food restaurants to adopt animal rights reforms. While there is a considerable body of research on the effect of consumer boycotts as an instrument to change the behavior of economic agents (Baron, 2003; Innes, 2006; Baron and Diermeier, 2007), and the ability of private activism to substitute for public regulation (Egorov and Harstad, 2017), the effect of market mediation in dampening the effect of private politics is less well-understood.

In many cases, market intermediaries stand between consumers and the target of consumer activism, for instance when a firm producing a boycotted product sells it through a series of distributors and retailers. If the intermediary has the ability to set prices, then adjusting relative prices for boycotted and non-boycotted products has the potential to effect the ultimate impact of private politics on economic agents. In this paper, we examine the role of market mediation in determining outcomes of a recent consumer boycott of microbeaded toothpaste products in the US.

Recent empirical work has documented the effectiveness of private politics in changing economic behavior. Politically-driven activism by consumer groups has been shown to result in decreased demand for French cars in China (Hong et al., 2011), French products in the US during the Iraq war (Pandya and Venkatesan, 2016), and imports of Danish products by Muslim countries following publication of the Jyllands-Posten Muhammad cartoons (Heilmann, 2016). For products with environmental health impacts, mercury warnings have been demonstrated to result in decreased consumer demand for pelagic fish species (Shimshack et al., 2007; Rausser et al., 2009), and advisory labels on sustainable fish stocks have been shown to shift demand from threatened species towards industries with sustainable practices (Hallstein and Villas-Boas, 2013). Yet, none of these studies consider the potential effect of market intermediaries in ameliorating the impact of private politics by adjusting relative prices.

We examine market mediation by retailers in the US toothpaste market in response to a consumer boycott of small, polyethylene particles known as ”microbeads” in personal care products. In 2014, a public outcry emerged over environmental contamination by microbeads, which adhere to toxic chemicals as they pass through fine mesh screens in sanitation systems and into waterways. Consumer awareness of the environmental health risks of microbeads was stimulated by an article in the Los Angeles Times on January 25, 2014 that documented plastic pollution in the Los Angeles River, sparking a boycott by California consumers on products containing microbeads. Over the following two years, several states introduced legislation to ban microbeads, which ultimately resulted in the Microbead-Free Waters Act of 2015 that phased out the use of microbeads in cosmetic products sold in the US in July 2017.

We examine microbeaded toothpaste sales at supermarkets in the ten states that introduced legislation to ban microbeads over the period 2014-2015. The toothpaste market represents an ideal market to study the effect of market mediation on private politics for several reasons. First, demand for toothpaste is relatively stable at the category level, with limited substitution possibilities available outside the category. Second, unlike private politics targeted at specific brands, the consumer boycott of microbeads affected several, major toothpaste product lines sold by leading toothpaste brands (Crest and Colgate), which also sell products without microbeads. These features of the data provide us with a unique, empirical opportunity to look for product line switching behavior within brands in response to private political action. Third, the toothpaste market allows us to examine how retailers mediate the effect of private politics by exploiting variation in how retailers respond to demand shocks at the individual product level while controlling for brand- and retailer-specific effects. Finally, supermarket data is syndicated by Nielsen, which allows us to make use of large sample panel data that contains considerable brand-level variation in boycott and non-boycott products within the category.

Consumer boycotts driven by a subset of consumers in a market can result in both an inward shift and a counter-clockwise rotation of market demand. For example, the subset of consumers participating in the boycott may value an affected environmental attribute that is not in the utility function of all consumers. In general, a boycott that persuades some but not all consumers to stop buying a product has the potential to alter the price elasticity of demand when the boycott alters the dispersion of consumer valuations in the market (Johnson and Myatt, 2006). Under circumstances in which a consumer boycott shifts a demand function inward, but makes it more price elastic, imperfectly competitive retailers have an incentive to narrow price-cost margins on boycotted products.[1] Indeed, the net effect of private politics can result in a rise in the total quantity sold in the market –a boycott backfire– in cases where retailers selectively reduce market prices on boycotted products

[1]Dubois et al. (2017) find evidence that potato chip advertising in the UK makes consumers less price sensitive, suggesting that a proposed ban on advertising would cause UK retailers to reduce potato chip prices.

relative to non-boycotted products.

Our analysis also contributes to a large literature on how retailers respond to contemporaneous demand shocks. Evidence from this literature generally suggests that retailers reduce prices and promote products more frequently at the category-level during periods of increased demand (Warner and Barsky, 1995; MacDonald, 2000; Chevalier et al., 2003). However, Nevo and Hatzitaskos (2006) provide evidence that much of the observed decrease in retail prices at the category level during periods of peak demand is driven by brand substitution from high-priced to low-priced products, a finding consistent with an influx of deal-oriented consumers. Our approach allows us to isolate product-level demand shocks among individual brands within the microbeaded and non-beaded segments of the toothpaste market, which provides a unique lens to examine how retailers adjust prices in response to demand shocks arising at the individual product level.

We frame our observations around two empirical approaches. First, we utilize retail scanner data in an event study that exploits variation across states in the timing of the boycott to derive inferences on demand changes and retail price responses. Specifically, we rely on the intensity of consumer Google searches for the term ”microbeads” to examine changes in the market share of microbeaded toothpaste products during and immediately after the initial boycott period. We divide this analysis into three parts: (i) we examine evidence that the boycott reduced demand for microbeaded toothpaste products by comparing changes in the market share of products in the beaded and non-beaded segments of the toothpaste category during periods of peak consumer search; (ii) we investigate retailer price adjustments at the product-level following the consumer boycott; and (iii) we then re-examine changes in the quantity of microbeaded toothpaste sales over the subsequent four week period after the peak in consumer search.

Second, we follow Johnson and Myatt (2006) in estimating changes in the dispersion of consumers' valuations for microbeaded toothpaste products during the boycott period. This allows us to estimate changes in demand elasticity across products in each segment of the market by exploiting variation in purchases by individual households in the Nielsen consumer panel during periods of peak search.

Our main findings are as follows. In response to the boycott, we demonstrate that consumers abruptly reduced their consumption of toothpaste containing microbeads.[2] Yet, in the period immediately following the boycott, we find that retailers reduced prices and promoted microbeaded toothpaste products more frequently relative to non-beaded products. As a result of these selective price discounts, we show that the total quantity of microbeaded toothpaste sold by retailers increased by 3-6 percentage points during the four-week period following spikes in consumer search. Our evidence therefore suggests that market mediation caused the consumer boycott to backfire.[3] We corroborate this finding by demonstrating that the consumer boycott decreased the dispersion of consumers' valuations for microbeaded toothpaste products relative to non-beaded products. The resulting increase in the relative price elasticity of beaded toothpaste demand provides explanation for our finding that retailers offered selective price discounts on microbeaded toothpaste products.

The remainder of the paper is organized as follows. In the next section, we provide background detail on the consumer boycott of microbeaded products. In Section 3, we

[2] Because microbeaded products vary both within and across brands, this finding is unlikely to be driven by changes in consumer purchasing patterns due to concurrent brand shocks.

[3] Moreover, we find microbeaded toothpaste sales to be more volatile during the post-boycott period than in the period prior to the boycott, which is consistent with our hypothesis that the boycott increased the price elasticity of demand in the microbead segment of the market.

develop a theoretical model to explain how private politics that shift demand inward for a boycotted product can result in a greater quantity of the boycotted product being sold. In Section 4, we develop our quasi-experimental approach and examine the change in consumer and retailer behavior in the period before and after the boycott. Section 5 formally estimates the change in dispersion of consumers' valuations following the boycott and presents evidence consistent with a relative increase in the price elasticity of demand among products in the microbeaded segment of the toothpaste market.

2 Background

Microbeads are solid, plastic polymers that are stable in response to physical and chemical triggers, causing them to persist in the environment. They are typically smaller than 200 μm, which allows them to pass through filtration screens in wastewater treatment systems, where they flocculate with other sewage particles in digestion tanks. The low density of microbeads (less than 1.0) allows them to float through storm water, sewage, and septic systems into tributaries, where they ultimately enter estuarine and marine ecosystems, resulting in environmental damage (Gregory, 1996; Browne et al., 2011).

In the US, plastic microbeads in personal care products are exempt as a drug or cosmetic under the Toxic Substances Control Act of 1976. As a consequence, these particles remained unregulated until passage of the Microbead-Free Waters Act in 2015. Prior to passage of this Act, the most common polymers used in toothpaste were polyethylene and polypropylene (see Table A1 for a list of microbead ingredients found in toothpaste).

Prior to 2015, identifying toothpaste with microbeads in a supermarket required consumers to recognize possible ingredients, notably "polyethylene," and verify whether or not any of these ingredients were present in a particular product. Although most toothpaste brands list ingredients on toothpaste tubes as well as on packaging, an ingredients list is not required on toothpaste tubes, making it difficult for consumers to verify microbead status of toothpaste at home.

At the time of the consumer boycott, virtually all toothpaste products containing plastic microbeads were contained in product lines sold by the leading toothpaste brands, Crest and Colgate (see Table A2 for a list of product lines containing microbeads in the Nielsen sample). News stories about the environmental risk of microbeads, which precipitated the consumer boycott, often failed to list the affected products, forcing consumers who wished to participate in the boycott to distinguish between "polyethylene," "polyethylene glycol," which is not a microbead, and "propylene glycol," a mineral oil, at the time of purchase.[4]

Environmental risks associated with microbeads were first acknowledged by the scientific community over twenty years ago (Gregory, 1996). Since that time, growing public awareness on the use of microbeads is evidenced by the mobilization of internet resources. On October 14, 2010, a wikipedia page was created for "Microbeads," and in 2012, an international coalition of non-governmental organizations, including the American organization 5Gyres, launched the "Beat the Microbead" campaign. Beginning in 2013, this campaign applied pressure on manufacturers to phase-out the use of microbeads in consumer products. In December 2013, a peer-reviewed article documented the effect of microbead pollution in the Great Lakes (Eriksen et al., 2013), followed by a widely-read Los Angeles Times article in January 2014 that documented microbead pollution in the Los Angeles River.

[4]Potentially in response to the microbead boycott, most brands now list polyethylene glycol as "PEG."

Following the initial news about microbead risks, consumer awareness of microbead pollution spread across states. Newspapers in California, New York, and states adjacent to the Great Lakes were among the first to carry the story, followed by a pattern of news and online media attention that varied over time and across states.

In response to growing public awareness of microbead pollution in US waterways, a number of states proposed legislation to regulate the use of microbeads in consumer products. The State of Illinois proposed the first microbead bill in the US on January 28, 2014, followed by the Microbead-Free Waters Act (Bill A08744) introduced by the New York Assembly. By 2016, a total of ten states had introduced legislation to ban the use of microbeads: California; Colorado; Connecticut; Illinois; Indiana; Maine; Maryland; New Jersey; New York; and Wisconsin. In each state, public attention on microbead pollution prompted large spikes in Google search intensity for the term "microbeads." Figure I shows the peaks in search activity for states that introduced legislation in 2014, where the peak search intensity almost always corresponds to the week that legislation was initially proposed in that state.

In the analysis to follow, we examine the effect of a consumer boycott of microbeads in the sample of ten states that adopted legislation to ban the use of microbeads in consumer products. Legislators in these states responded decisively to consumer pressure to eliminate microbeads, which suggests the impact of the consumer boycott on the demand for microbeaded products was likely largest during the period immediately preceding the ban in these states.

3 Theoretical Model of Intermediation

We examine the effect of a change in demand for microbeaded toothpaste in a quantity-setting oligopoly market.[5] Let $p(Q;\theta)$ denote inverse demand in the market, with $p_Q(Q;\theta) < 0$ and $p_\theta(Q;\theta) < 0$, where θ is a shift parameter that reflects the effect of the consumer boycott on market demand.

Retailer $i = 1, 2, ..., n$ selects the quantity q_i, resulting in retail profits of

$$\pi_i = p(Q;\theta)q_i - c(q_i).$$

where market quantity is $Q = \sum_{i \epsilon n} q_i$. The first-order condition is

$$\frac{\partial \pi_i}{\partial q_i} = p(Q;\theta) + p_Q(Q;\theta)q_i - c'(q_i) = 0.$$

We assume demand is downward-sloping, $p_Q(Q;\theta) < 0$, marginal cost is upward-sloping, $c''(q_i) > 0$ and that the oligopoly equilibrium satisifes the usual stability condition $p_Q(Q;\theta) + p_{QQ}(Q;\theta)q_i < 0$. The symmetric market equilibrium, $q^* = q_i$ for all i, is completely characterized as the solution to

$$p(nq^*;\theta) + p_Q(nq^*;\theta)q^* - c'(q^*) = 0. \tag{1}$$

Making use of the implicit function theorem on equation (1), the effect of a demand shift

[5] Under reasonably general demand conditions, Johnson and Myatt (2015) show that the multi-product Cournot outcome in a quality-differentiated product category results in equilibrium prices that align with the prices set by retailers in a single-product Cournot market.

on the quantity sold by a retailer is

$$\frac{\partial q^*}{\partial \theta} = \frac{-\left(p_\theta + p_{Q\theta} q^*\right)}{(n+1)p_Q + p_{QQ} q^* - c''},$$

where the denominator is negative by the stability condition. It follows immediately that a consumer boycott associated with a negative shock in demand, $p_\theta < 0$, increases the quantity sold by retailers whenever $p_{Q\theta} q^* > -p_\theta$. When demand rotates counter-clockwise as a result of a demand shock, $p_{Q\theta} > 0$, retailers select narrower price-cost margins, resulting in lower retail prices for the boycotted product. If the rotation effect is sufficiently strong, the increased quantity demanded by non-boycotting customers at lower prices is more than enough to compensate for the decrease in demand by boycotting customers.

4 Data

To examine how the consumer boycott on microbeads affected the toothpaste market, we rely on four types of information: (i) weekly retail data; (ii) information on toothpaste ingredients; (iii) a measure of consumer awareness of risks from microbeads; and (iv) county characteristics.

We measure consumer responses to news of the environmental risks from microbeads using weekly toothpaste data for ten states supplied by Nielsen through the Kilts Center for Marketing. Nationally, Nielsen data cover about half of all food and drug retailers and about a third of mass merchandisers.[6]

[6] The conclusions drawn from our study are ours alone and do not reflect the views of the Nielsen Company, LLC. Nielsen is not responsible for, had no role in, and was not involved in analyzing and preparing the results reported herein.

We rely on three different methods of aggregation to investigate brand market share, retailer pricing, and total sales in the beaded and non-beaded segments of the market following the boycott. In each case, we control for changes in demand and retailer pricing around the first of each month due to access to entitlement programs and paychecks (Hastings and Washington, 2010).

Toothpaste brands offer multiple product lines, and within these lines, provide a variety of products from which consumers can choose. For example, Crest offers product lines such as Pro-Health and 3D White. Individual products within each product line are identified by retailers with a unique universal product code (UPC), for instance a 5.8 oz tube of Crest 3D White with extreme mint flavor. We expect a consumer boycott of microbeads to affect all toothpaste products within product lines containing microbeads, while having a negligible effect on product lines that do not contain microbeads.

Market Share Growth. To identify the effect of a boycott on microbeaded products, we follow Pandya and Venkatesan (2016) by focusing on the change in the annual growth rate of market share for each product line in a store. Specifically, we use market share for each product line to estimate the effect of the microbead boycott during the boycott week relative to market share during the same week of the previous year. Because stores may vary in aggregate demand within the product category, examining market shares allows us to scale changes in demand for a product line to store-level demand, thereby allowing us to identify shifts in demand across individual product lines within brands in a store.

We use annual changes in market share to difference out time-invariant store characteristics and to control for seasonal fluctuations by store and by week. Because the growth rate in market share may be non-stationary, differencing the growth rate enables consistent estimation with ordinary least squares.

Suppose product line i has h products in store k in week t in 2014. We define total units sold across products for line i in store k as $units14_{ikt}$ and total units sold within the entire product category in store k as $units14_{kt}$, so that a product line's market share in store k in week t is given by $share14_{ikt} = \dfrac{units14_{it}}{units14_{kt}}$.

We examine the annual difference in market share for each product line in a store. For a comparison of market shares for a given week between 2013 and 2014, we define the annual change in weekly shares to be $share1314_{ikt} = share14_{ikt} - share13_{ikt}$. Examining changes in market share growth for similar weeks within a given year controls for seasonal patterns that can emerge at the store and product-line level, which allows us to identify changes in the growth rate of market share around boycott events. We examine changes in the weekly growth rate of market share for product line i at store k at time t in response to shocks measured on a weekly duration as $\Delta share1413_{ikt} = share1413_{ikt} - share1413_{ik,t-1}$. We confine our analysis to product line-stores that have positive sales within the category during each week of 2014.

The resulting distribution of the growth of market share is symmetric around zero. Figure II shows the distribution, which has a mean of -0.0001 and a standard deviation of 0.059.[7] The panel begins the first week of January 2014 and ends the week of the boycott, which varies by state. Our panel restriction, that the product line have some sales every week, results in a total of 56 product lines, six of which contain microbeads while the remaining products do not.

[7] Comparing these figures to Pandya and Venkatesan (2016), who aggregate products to the brand level, the standard deviation in our sample is slightly smaller while the mean is much larger.

Retail Pricing. To investigate the retailer's pricing response, we choose our unit of observation to be a UPC-retailer-state. To do so, we calculate the unit-weighted average price per ounce for each UPC sold by a given retailer in a given state. The average price in our sample is $0.785 per ounce, with a standard deviation of $0.53. Figure III shows the distribution of the price per ounce, which exhibits a long tail toward higher prices, as one might expect. This feature of the data is consistent with the product-level regressions of Nevo and Hatzitaskos (2006).[8] In addition to the price per ounce, we define goods as "On Sale" if their price is below the mode price within a six week window as in Chahrour (2011).

An alternative approach of examining retail pricing behavior at the UPC level would be to create a price index for toothpastes of different size and flavors within a store in the spirit of Chevalier et al. (2003). However, as Nevo and Hatzitaskos (2006) observe, such a price index masks changes in the composition of products sold that are unrelated to changes in retail prices, for instance if consumers substitute from high-priced to low-priced products during periods of peak demand, confounding inferences at the individual product level. Another alternative would be to select products sold each week to create a balanced panel; however, this causes problems if the selected product-stores are more price inelastic than those not selected, as demonstrated by Gandhi et al. (2014). The panel begins the first week of January 2014 and ends four weeks after the boycott, the timing of which varies by state.

Quantity Sold. We estimate the effect of the microbead boycott on total ounces sold by retailers for toothpaste with and without microbeads. We measure the ounces sold of a particular product based on the size, number of items per pack, and total number of units

[8] Aggregating at the UPC-retailer-state is appropriate, because most retailers set nearly-uniform prices across their stores (DellaVigna and Gentzkow, 2017).

sold for a product line in each retail store in each week. We aggregate across product lines so that for each store-week, we have two observations: total ounces sold of products with beads and total ounces sold of products without beads. This measure ensures a nearly balanced panel and while holistically capturing pollution from microbeads. Across these categories, the average number of ounces sold per week is 252 oz, with a standard deviation of 312 oz. To accommodate variation in store-level demand across stores, we take the natural log of ounces sold.[9] Figure IV shows the distribution of the natural log of ounces sold, which is well-behaved. Like in the case of retail prices, the panel begins the first week of January 2014 and ends four weeks after the first boycott.

Microbeads in Toothpaste. Nielsen does not report product ingredients. As a result, we rely on several sources to identify product lines and flavors that contain microbeads. These sources include lists developed by non-governmental organizations Californians Against Waste, 5Gyres, the Household Products Ingredient Database,[10] which we supplement with information collected from manufacturer websites, blog posts, and images of product labels posted on amazon.com. The list of ingredients classified as microbeads appears in Table A1 and the list of product lines and flavors with beads is contained in Table A2. Based on this information, we match product lines containing microbeads in the Nielsen database. Because a small number of product lines contain microbeads in some, but not all flavors, we code product lines in which bead content varies by flavor as having microbeads for our analysis of market share growth at the product line level.[11]

Consumer Awareness. Our preferred measure of consumer awareness of microbead

[9]Because we have some weeks with have zero sales, we shift the distribution and use the natural log of the ounces sold plus one.

[10]https://householdproducts.nlm.nih.gov/index.htm, accessed April 2016.

[11]An exception is Colgate's Icy Blast flavor, which is a small share within the main Colgate brand.

pollution is derived from Google Trends. We rely on Google Trends to independently assess weekly search frequency for the term "microbeads" among consumers in each US state. We focus on the ten states that introduced legislation prior to the National ban on microbeads in 2017: California. Figure I provides details on Google Trend scores for 2014 by state.

Our Google Trends search period is from January 6, 2013 to February 6, 2016.[12] Google Trends measures search intensity by pulling a random sample of search queries from the period of inquiry. For each week, the total number of searches for the term "microbeads" is divided by the total searches in that state during that week to provide an index of the relative popularity of the search among all Google searches during that week. These values are then scaled relative to the week with peak "microbead" searches by assigning the week with the highest proportion of Google searches an index value of 100.

We choose a relatively long period of inquiry with Google Trends to identify the most important spikes in search intensity from the beginning of the period of consumer microbead awareness.[13] To identify the first peak in consumer awareness, we use a threshold score of 20 for search intensity.[14]

County Characteristics. In our robustness tests, we break out the boycott and backfire responses across county characteristics. To capture county sensitivity to environmental issues, we use the lifetime score for representatives from the 2015 League of Conservation Voters environmental scorecard (LCV, 2015). Since districts cross county lines, we take the average score for all congresspeople that represent any part of the county. The median score

[12] We conducted our Google Trend inquiries on May 10, 2016

[13] A shorter period of inquiry, e.g. examining search intensity for only 2014, would peg one week of 2014 as having a score of 100.

[14] Our results are robust to the selection of different thresholds. The precision of the estimated boycott coefficient is unchanged when varying the threshold and the results are qualitatively similar in each case.

for the counties in our ten state sample is 77. We use a county score of greater than 75 to denote "High Green" counties, meaning those with legislators who more often vote for environmental protection. To check whether the backfire response is driven by stockpiling, we compare the response across county median household income in 2014 from the U.S. Census Bureau, Small Area Income and Poverty Estimates (SAIPE, 2015). The median income across counties is $56,645. In our robustness check, we bin households into low, middle, and high income, which are defined as less than $50,000, between $50,000 and $75,000, and over $75,000. These values are close to incomes at the 25th and 75th percentiles.

5 Empirical Framework

To identify the impact of the consumer boycott on the retail market for microbeaded toothpaste products, we exploit both the variation in microbead use across product lines within a brand and the geographic variation in how consumer awareness of microbead risk disseminated across states in our sample. We first test for whether consumers boycott. Next, we test for whether retailers respond to demand shocks in the microbeaded segment of the toothpaste market by offering selective price discounts on products containing microbeads. Then, after allowing for a period of price adjustment by retailers, we return to examine the effect of the consumer boycott on the sales velocity of microbeaded products during the four-week period following the initial spike in consumer search intensity.

Our identification strategy is to compare changes in outcomes for product lines with microbeads relative to products lines without microbeads during periods before and after the boycott. To do this, we employ a difference-in-differences estimation strategy using a fixed effects regression framework. In our baseline specification, the identifying assumption is parallel trends between beaded and non-beaded product lines prior to the boycott. We relax this identification assumption in two other specifications. For our specifications that include brand trends, the identifying assumption is that there is no discontinuous change in trend for beaded products at the point of time of the boycott apart from the boycott event itself. For our specifications that include brand interactions, the identifying assumption is that there are parallel trends between beaded and non-beaded product lines within a brand prior to the boycott

5.1 Change in Share Growth

We specify our regression on the change in the growth rate of market shares as follows. For product line i of brand j in store k in state s in week t, we estimate a difference-in-differences model of weekly changes in each brand's rate of market share growth:

$$\begin{aligned} \Delta Share14-13_{ijkst} = \alpha_{ik} &+ \beta_0 HighGoogle_{st} + \beta_1[Bead_i * HighGoogle_{st}] \\ &+ \beta_3 \Delta Price14-13_{ijkt} + \beta_4 \Delta NumVariants14-13_{ijkt} \\ &+ \beta_5 \Delta Share14-13_{ijkt-1} + \epsilon_{ijkt-1} \end{aligned}$$

where α_{ik} are fixed effects at the product line-store level, and $Bead_i$ is bead status. Within the product mix of each brand, we identify whether product lines with microbeads respond differently to spikes in search intensity relative to product lines without microbeads by interacting high search intensity events in each state with bead status.

Our estimation approach follows Pandya and Venkatesan (2016). However, instead of using a random effects framework, we estimate our model using a fixed effects framework.[15] Like Pandya and Venkatesan (2016), we cluster standard errors at the store level. We include an additional control for the first week of the month in all regressions to account for changes in pricing and demand that week (Hastings and Washington, 2010). We use two strategies to assess the robustness of our results. First, we stratify the sample by environmentalism. If market share growth decreases due to a boycott, we would expect the effect to be concentrated in "High Green" counties. Next, we include differential trends for products with and without microbeads.

Our goal is to investigate changes in consumer demand during periods of peak search intensity, which mark the beginning of the boycott period. The timing of the treatment variable, $HighGoogle_{st}$, and the length of the panel vary across states in our analysis. For the boycott period, $HighGoogle_{st}$ is a dummy variable equal to one the first week of 2014 that search intensity spikes (above the index value of 20) for consumers in a given state. The panel begins in the first week of January 2014 and ends at the conclusion of the first week of peak search intensity in each state.[16] The coefficient on $HighGoogle_{st}$ thus reflects the change in the rate of market share growth in the first week of the boycott relative to the weeks prior to the boycott in that year.

5.2 Retail Price Adjustment

To investigate retailers' price response to the boycott, we estimate price changes following a surge in microbead search intensity at the product-retailer-week level. Our analysis is framed

[15] Pandya and Venkatesan (2016) use this estimation approach as a robustness check.

[16] We truncate our panel at this point to make a clear distinction between consumers' awareness of microbead risk (the boycott) and retailers' response to the resulting demand shock in the microbead segment of the toothpaste market.

by a panel that runs from January 2014 to four weeks after the first high search intensity event in each state. Because our goal is to identify changes in retail pricing in response to a consumer boycott, and retailer responses are not instantaneous, we examine retail prices over the entire four week period following a spike in search intensity as opposed to terminating our sample after the initial boycott week.

Our retail pricing specification follows Nevo and Hatzitaskos (2006). For product h of product line i of brand j sold by retailer r in state s in week t, we estimate a triple difference-in-differences ordinary least squares model of the retailer's average price that week:

$$\begin{aligned} price_{hijrst} = {} & \alpha_{hijrs} + \beta_0 HighGoogle_{st} + \beta_1 [Bead_i * HighGoogle_{st}] \\ & + \beta_2 [Brand_j * HighGoogle_{st}] + \beta_4 Post_{st} + \beta_5 [Bead_i * Post_{st}] \\ & + \beta_6 [Brand_j * Post_{st}] + \delta_t + \epsilon_{hijrst} \end{aligned}$$

where $price_{hijrst}$ is either the sales-weighted average price per ounce or an indicator variable equal to one if the good is on sale. $HighGoogle_{st}$ indicates that the Google Trends score is at or above 20 and $Post_{st}$ indicates the four week period following the initial surge in search intensity. We include week of year fixed effects δ_t in specifications without trends. Because prices are aggregated across stores within retailer within a state, we expect correlation among products within a retailer within a state and, accordingly, cluster standard errors by retailer-state.

5.3 Quantity of Toothpaste Sold

To investigate the possibility of a boycott backfire, we examine changes in the total quantity of beaded toothpaste sold by retailers in the four weeks following the boycott. The equilibrium quantity of beaded toothpaste sold by retailers is the product of two effects: (i) reduced consumer demand for microbeaded products following the boycott; and (ii) the change in quantity demanded resulting from retail price adjustment over beaded and non-beaded product lines. We refer to cases in which the sales velocity of microbeaded products rises following the consumer boycott as a boycott backfire.

We specify the quantity regression identical to our pricing regression with the exception that our unit of observation differs. Instead of aggregating our data to the level of a retailer-product-week, we aggregate our data to bead-store-week. This ensures a nearly balanced panel and holistically captures pollution via changes in total ounces sold of bead products.

For the backfire regression, $Post_{st}$ refers to the four weeks that follow a period of high search intensity. The coefficient thus reflects the percentage change in the total quantity of sales for the four weeks following the consumer boycott relative to the weeks before the boycott. The panel length is identical to the panel in the retail pricing regression and we cluster standard errors at the store-level as we did in the boycott regression.

5.4 Identifying Assumptions

We assume the timing of each spike in Google search intensity is independent of changes in state-level product line trends. This assumption would be violated if consumers are compelled to search for information on microbeads during times that happen to coincide with periods of falling market share for beaded product lines. While such an outcome is possible, for instance if newspapers tend to run stories about microbead pollution during periods of softening demand for product lines with microbeads, we believe it is unlikely for several reasons. First, the rate of share growth would need to differ across states due to differences in the timing of searches across states, which varies considerably in our sample. Second, the primary trigger for state-level spikes in search intensity appears to be local newspaper reporting in that particular state. Reverse causality would require local newspapers to be both aware of the rate of market share growth of microbeaded products and responsive to it. Finally, raw data on market shares across states reveals that the overall market share of microbeaded products was trending upwards at the time of the initial boycott in each state.

To separate consumers' reactions to the microbead boycott from retailers response to the resulting demand shock, we assume the duration of the consumer boycott is consistent with the observed spike in search intensity for the term "microbeads" above a tolerance level of 20 on our index. This is consistent with a large body of marketing research that demonstrates consumers' responses to television advertising is short-lived, typically lasting only one to two weeks (Gerber et al., 2011).

6 Results

We begin by investigating the immediate response to news about microbead pollution. After characterizing this initial impact of the consumer boycott, we turn to retailers' price responses over the four weeks following the spike in consumer awareness. Finally, we examine the net effect of the boycott on the consumption of products with microbeads over the four week period following the demand shock, after controlling for both the change in demand and the retail pricing response.

6.1 Sales Velocity Initially Falls

Table 2 presents results from our fixed-effects OLS regression on changes in the growth of market share for a product line within a store. Moving across columns in the table introduces robustness checks to the basic specification. The "High Google" coefficient reflects the average change in the annual market share growth rate for toothpaste product lines during the initial week of intensive consumer search within that state. Our primary focus is on the interaction between High Google and Bead, which reflects the change in the growth rate of market share for product lines that contain microbeads relative to non-beaded products.

The coefficient on the interaction of High Google and Bead is negative and statistically significant at the 1 percent level.[17] The rate of market share growth for product lines containing microbeads declines by 1.3 percentage points, a remarkably strong effect given that consumers needed to identify products containing microbeads from among a large family of Crest and Colgate brands that did not explicitly label products as microbead-free.

It is useful to compare the magnitude of the boycott effect to that of Pandya and Venkatesan (2016) (PV). Our specification differs from PV in two ways that may affect the magnitude of our coefficient estimates. First, we exploit geographical variation across states in the tim-

[17] Caution should be exercised in interpreting this result. If consumers shift from beaded to non-beaded products within a given brand, then non-beaded products will gain market share within brand at the expense of beaded products, over-stating the magnitude of the treatment effect on microbeaded products. To test for potential violation of our parallel trends assumption, we run a validity check on the model based on a difference-in-differences comparison between Illinois and Wisconsin, which have boycotts at different times. This analysis results in a statistically significant coefficient estimate for the decrease in market share for microbeaded products.

ing of news reports on microbead risk, whereas PV examine information shocks from Fox News, which is nationally syndicated. Second, we focus on one product category (toothpaste) instead of a large group of product categories, which allows us to control for time and brand-specific fixed effects. In PV, the maximum predicted change in rate of market share growth for "French-sounding" products in response to anti-France comments was a decline of about 0.36 percentage points, approximately one-quarter our estimated magnitude.

To examine whether the effect is driven by environmental boycotts, we stratify the sample into Low Green counties, where the average lifetime environmental score for elected representatives is at or below the 75, and High Green counties with a score above 75. If the decrease in market share growth in the High Google period is due to a boycott, we would expect to find little change in the Low Green counties and a larger change in High Green counties. Comparing columns two and three, this is what we find. The coefficient is larger and more precise for High Green counties; market share declines by 2.7 percentage points. In Low Green counties, we fail to find an effect.

Our estimated boycott effect is robust to additional controls, including a triple difference specification and time trends by beaded products. While all specifications control for product line-store seasonal changes through annual differencing, the specification in column (5) includes trends for bead and non-bead products, allowing these products to have different trends prior to the High Google event. The coefficient on the High Google and Bead interaction fails to change with application of these additional controls. In all specifications, we include controls for the change in the growth rate of the average price and the number of individual products offered in the product category. Like PV, we find higher prices to be negatively associated with annual growth in market share and product variety to be positively associated with annual growth in market share.

6.2 Retailers Respond with Selective Price Discounts

Table 3 presents results from our retail pricing model that estimates changes in the unit price per ounce at the UPC-level by state and retail chain. The *Post* coefficient reflects the change in average price in the four weeks following the initial spike in search intensity in each state. The panel begins January 2014 and terminates four weeks after the spike.

Our primary focus is on the interaction term on microbeaded toothpaste products. Notice that the average price for beaded products falls by 0.85 percentage points in the four-week period following a high search intensity event. Prior to the boycott, prices for beaded products were generally rising; consequently, including trends in column two increases the coefficient estimate to 1 percentage point.

Recall that the post boycott period covers four weeks. Therefore, the interpretation of the $Post * Bead$ coefficient is the average decrease in the retail price of microbeaded toothpaste products relative to non-beaded products across the entire four-week period. The entries in Table 3 suggest that retailers respond to the consumer boycott of microbeads by offering selective price discounts on microbeaded toothpaste products.

It should be noted that our specification cannot distinguish between a temporary price decrease (i.e., a promotion on microbeaded products) or a lower modal price. However, in general, sales account for approximately 20-50% of price variation in conventional retail stores (Hosken and Reiffen, 2004), and the average duration of sales for processed and unprocessed groceries is about two weeks (Nakamura and Steinsson, 2008). Using the Dominick's dataset for Chicago, Levy et al. (2011) found that the average price reduction for toothpaste was between \$0.35 and \$0.42 cents per unit during periods of price promotion. Based on the average toothpaste price of \$0.80 per ounce and the average package size of 5 ounces in our sample, our estimate of an additional 1% price decrease for beaded toothpaste products is consistent with an average price decrease across the four weeks of \$0.04.

As an alternative to the price per ounce, columns three and four report coefficient estimates where the outcome is a binary variable equal for whether the product is on sale. The entries in these columns provide additional evidence of selective discounts for microbead products: the likelihood a microbead product being on sale in the post-boycott period increases by 4 percentage points. The estimate is stable when including trends and statistically significant at the 1 percent level in both cases.

6.3 Boycott Backfire

Table 4 shows the effect of the boycott on the quantity of microbeaded products sold by retailers in the four weeks following the initial spike in search intensity. The results are derived from an OLS regression with fixed effects in which the outcome variable is the natural log of the total ounces sold for bead and non-bead products in a given retail store. As in Table 3, the panel begins January 2014 and ends four weeks after the first spike in search intensity.

The $HighGoogle$ coefficient captures the average change in ounces sold during the week(s) in which search intensity exceeds the threshold of 20 in our index. As in the case of the market share growth regression, the $HighGoogle * Bead$ coefficient is negative and statistically significant in all specifications at the 0.1 percent level. The initial sales velocity on product lines containing microbeads declines significantly relative to products that do not contain microbeads. Columns two, three, and four stratify the sample by county median income. We find evidence of boycotting behavior in each type of county.

Our findings corroborate our earlier boycott regression results based on growth in market shares. The initial effect of the boycott was to reduce demand for products containing microbeads. Moreover, while our earlier estimation of changes in the rate of market share growth obscure aggregate demand effects, the entries in Table 4 reveal that the average velocity of microbead toothpaste sales decreased by about 5.7 percentage points. Unlike findings in the literature during periods of product recalls for infant toys (Freedman et al., 2012) and Italian cheese (De Paola and Scoppa, 2013), we fail to find a spillover in negative demand to other manufacturers.

Our main coefficient of interest is the interaction term $Post * Bead$, which reflects the differential change in retail sales velocity for product lines containing microbeads relative to the average change in sales velocity for lines without microbeads. Relative to the period before the boycott, sales velocity significantly increases for toothpaste products containing microbeads during the four-week period after the initial spike in search intensity.

The coefficient $Post$ reflects the average change in quantity of all toothpaste products during the four weeks following a boycott. The average quantity of toothpaste sold by retailers in the post period is unchanged relative to the pre-boycott period. In contrast, retail sales velocity increased for toothpaste products containing microbeads in the four-week period following the boycott, resulting in a greater quantity of microbeaded toothpaste being sold. Indeed, retailers increased sales velocity for microbeaded toothpaste products by over 3 percentage points relative to non-beaded products. Our evidence indicates the boycott backfired: Retailers sold a greater quantity of microbeaded toothpaste products due to selective price discounting behavior in the period immediately following the boycott.

The increase in microbeaded toothpaste consumption can be explained in light of our earlier retail pricing results in Table 3. Over the four week post-boycott period, retailers selectively decreased prices on microbeaded products by an average of $0.04. In response to the selective price discount, it is possible that existing consumers stockpiled microbeaded toothpaste products, which would tend to exaggerate our estimate of the boycott backfire. Nevertheless, because stockpiling behavior is sensitive to the duration of a temporary price decrease, the longer the duration of retail price reduction, the weaker the resulting stockpiling effect (Hendel and Nevo, 2006). Furthermore, we find evidence of increased sales for low income counties where consumers are less able to stockpile (Orhun and Palazzolo, 2018).

Overall, we find evidence that retailers selectively applied a large and persistent price discount on microbeaded toothpaste products over the period immediately following the consumer boycott. The selective price discount induced a commensurately large and persistent increase in sales velocity, as shown in Figure II, an outcome that is difficult to explain by stockpiling behavior alone. Persistent price decreases can induce consumers to switch brands, particularly for products like toothpaste that are relatively inconspicuous (Bronnenberg et al., 2012). In this regard, the boycott has dynamic implications; it could increase the set of consumers loyal to product lines with microbeads, further increasing sales velocity over time.[18]

[18] We conduct a t-test using our household panel on whether there were more frequent first-time buyers for beaded products in the period following a high Google event. Based on this test, we reject the null hypothesis that the frequency of first-time buyers was the same for beaded and non-beaded toothpaste at the .01% level for Crest, although we fail to reject it for Colgate.

6.4 Robustness

So far we have shown an initial decrease in consumption of microbeaded toothpaste in response to the consumer boycott in a panel difference-in-differences framework both with market share growth and ounces sold as the outcome variable. Here, we address three threats to identification: (i) potential violation of parallel trends; and (ii) whether the uptick in Google search activity anticipates the boycott as opposed to the boycott prompting search.

Our identifying assumption is parallel trends for beaded and non-beaded toothpaste products prior to the spike in consumer awareness of microbead risk. In each table, we relax this assumption in the final column by allowing trends by bead status. Here, we further test our assumption of parallel trends using an event study approach. These results are reported in Figure II, which is constructed by fitting a local polynomial to the natural log of ounces sold of products with and without microbeads, across states, over a duration of four months before to two months after the initial boycott date. The data used to estimate the model are the same as in Table 4 except the observations were de-meaned. Notice that the trends in market share are similar across the two categories before the event and only diverge afterwards. The increase in sales for beaded products is about 4 percentage points, which is fully sustained over the ten weeks following the boycott. This evidence suggests that the price decrease caused at least some consumers to switch from non-beaded to beaded toothpaste products.

We test for sensitivity of our identification strategy to reverse causality. If demand happens to decrease for a product at exactly the time that people begin Googling for more information about it, then the uptick in Google search activity would coincide spuriously with the change in demand, independent of a consumer boycott on microbeads. However, Figure II suggests that changes in market share did not precede the increase in Google search activity.[19]

As an additional robustness check on our results, we conduct an alternative analysis under a different identifying assumption. Because consumer search intensity varies both across states and over time, we examine sales of microbeaded toothpaste products in two neighboring states that introduced microbead legislation at different times, Illinois and Wisconsin. Here our control group is microbead toothpaste in a neighboring state instead of toothpaste products without microbeads. We assume parallel trends for microbeaded toothpaste sales in Wisconsin and Illinois, and examine the difference in market share growth between the two states at the time of the spike in search intensity in Illinois. Table A3 in the online appendix reports the change in market share growth for microbeaded toothpaste in Illinois, relative to Wisconsin during the period of peak Google search intensity in Illinois. Our estimate from this approach is similar in precision but much greater magnitude as the coefficient reported in Table 2, which suggests our main results are robust to this alternative control.

7 Structural Model

The results from our quasi-experimental framework provide evidence that retailers responded to a negative demand shock for microbeaded toothpaste products by selectively reducing retail prices on beaded product lines. Within four weeks of the spike in search intensity, retailers were selling a larger quantity of microbeaded toothpaste products than they sold in

[19] We also test whether the high Google dummy was a leading indicator of changes in market share and fail to find such an effect. These results are available upon request.

the period prior to the consumer boycott, resulting in a boycott backfire. In this section we develop and estimate a structural model to examine whether these observed outcomes can be explained by an increase in the price elasticity of demand for microbeaded toothpaste products relative to non-beaded products during periods of high Google search intensity.

We frame our structural model in terms of the underlying distribution of consumer valuations for microbeaded toothpaste products. Changes in the dispersion of consumers' valuations during periods of increased search intensity result in changes in the price elasticity of consumer demand facing retailers (Johnson and Myatt, 2006). We test for rotation effects in the market demand for microbeaded toothpaste using a discrete choice model of household demand that admits heterogeneity in consumer preferences over the boycotted attribute. Thus, we allow the market response to the consumer boycott to be randomly distributed over households, and allow shocks in Google search intensity to affect the dispersion of consumers' valuations for microbeaded toothpaste products. Our approach allows us to identify counter-clockwise rotations in demand ("elastic rotations") under circumstances where the dispersion in consumers' valuations decreases in response to the consumer boycott.

Based on our empirical finding that retailers mediated the effect of a consumer boycott on microbeads by reducing retail prices on microbeaded toothpaste products, we expect to observe the following structural demand effects. First, the consumer boycott should result in decreased market demand for microbeaded toothpaste products, as consumers boycotting microbeads exited the market. Second, in response to information shocks regarding microbead risks, the departure of boycotting types from the market for microbeaded toothpaste should reduce the dispersion of consumers' valuations among the remaining population of consumers. That is, we expect to observe a counter-clockwise rotation of the demand curve for microbeaded toothpaste.

7.1 Panel Data: Retail Toothpaste Purchases

We use the Nielsen consumer panel over the period spanning from January 2014 to one week after the consumers boycott in each of the ten states that ultimately proposed microbead legislation. Because households can make multiple trips on the same day and purchase the same product line on each trip, we average the product line price across households making more than one trip per day to a retailer. We drop duplicate observations from the analysis, so that each remaining observation can be interpreted as a household-product line-day.

We define the retail price as the net price per ounce paid by a household after coupons. For toothpaste products that contain microbeads, the average price per ounce is \$0.62, with a standard deviation of \$0.49. Summary statistics are reported in Table A4.

7.2 Econometric Model

Consider a unit mass of consumers, each willing to pay w for one unit of the product. The distribution of w is represented by $F_s(w)$, which is twice continuously differentiable in both s and w with density $f_s(w)$. The parameter s governs the shape of the distribution of valuations such that an increase in s represents a spread in the density of w and, hence a clockwise rotation of $F_s(w)$ about some point $\hat{f}$.

Now consider the effect of a spread in consumer valuations on the distribution of market demand. At any price, p, the proportion of consumers who purchase the good is given by: $q = 1 - F_s(p)$. Inverting this expression gives an expression for the inverse demand curve:

$P_s(q) = F_s^{-1}(1-q)$, so a change in s rotates the inverse demand curve in a manner analogous to the change in the distribution of consumer valuations.

To empirically measure the effect of a consumer boycotts on market demand for microbeaded toothpaste products, we estimate willingness to pay (WTP) at the household level using a random utility framework in which the distribution of consumer heterogeneity reflects the distribution of marginal valuations described above. A random utility model is appropriate for our data, because we analyze household-choices among differentiated consumer product goods. Choices on each purchase occasion, therefore, are necessarily discrete.

In the random utility model, consumer utility is the sum of deterministic and stochastic parts such that $U_{ij} = V_{ij} + \varepsilon_{ij}$ for product j by consumer i, where V_{ij} is the deterministic component of utility, and ε_{ij} is an iid error term. The deterministic component, in turn, is a function of attributes of the choice (z_j), whether a particular toothpaste is beaded (b_j), a vector of boycott events (g_i), and income (y_i) (Anderson et al., 1989).

Assuming only two products for clarity of exposition, the marginal value consumer i places on product $j = 1$ is defined as the amount of income that leaves her utility at least as great with or without the purchase:

$$V_{i0}(z_0, b_0, g_i, y_i) + \varepsilon_{i0} \leq V_{i1}(z_1, b_1, g_i, y_i - c_{i1}) + \varepsilon_{i1},$$

where c_{i1} is the marginal value of product 1 by consumer i. We solve for the willingness to pay by consumer i by invoking the random utility assumption and recognizing that $Pr(WTP_{i1} \geq c_{i1}) = \Pr(V_{i0} + \varepsilon_{i0} \leq V_{i1} + \varepsilon_{i1})$.

Assuming the error term is double-exponential distributed with mean 0 and variance

$\pi^2\mu^2/3$, consumer willingness to pay becomes:

$$Pr(WTP_{i1} \geq c_{i1}) = \frac{\exp(V_{i1}/\mu)}{\exp(V_{i1}/\mu) + \exp(V_{i0}/\mu)},$$

where μ is the logit scale parameter.

Solving for the willingness to pay from this expression, we write the odds ratio of choosing product 1 relative to product 0 as:

$$\frac{\Pr(j=1)}{1-\Pr(j=1)} = \frac{\exp(V_{i1}/\mu)/(\exp(V_{i1}/\mu) + \exp(V_{i0}/\mu))}{\exp(V_{i0}/\mu)/(\exp(V_{i1}/\mu) + \exp(V_{i0}/\mu))} = \exp(V_{i1}/\mu),$$

where $\Pr(j = 1)$ is the probability of purchasing good 1 and where we have normalized the utility of product 0 to unity. Taking logs of both sides of the odds ratio gives an expression for the willingness to pay by consumer i as a function of choice attributes, the boycott indicator and the scale parameter (which we normalize to 1 without loss of generality):

$$ln\left(\frac{\Pr(j=1)}{1-\Pr(j=1)}\right) = WTP_{i1} = V_{i1}/\mu.$$

With an appropriate specification for V_{i1} this modeling structure allows us to test for both the direct effect of the boycott on household willingness to pay for microbeaded toothpaste and the indirect effect of the boycott on changing the dispersion of consumers' valuations.

We consider a utility function that is additive over product attributes. Specifically, writing V_{i1} in terms of an estimable model of utility, we specify the following random utility model for person i consuming product j:

$$V_{ij}(z_j, b_j, g_i, y_i) = \tau_{ij} + \alpha y_i + \gamma_m + \sum_{k=1}^{K} \gamma_k m_{jk} + \delta_1 g_i + \delta_2 g_i b_j + \xi_j, \tag{2}$$

where τ_{ij} is a household-product-line-specific constant, α is the marginal utility of income, β_k are marginal values for each product attribute, γ_k represents the influence of each demographic attribute m_{kj} on willingness to pay, g_i is our measure of boycott-media intensity observed by consumer i, b_j is a binary indicator of whether a particular toothpaste is beaded ($b_j = 1$) or non-beaded ($b_j = 0$), δ_1 is the impact of a boycott shock on indirect utility for consumers of non-beaded toothpaste, δ_2 is the impact of a boycott shock on consumers of beaded toothpaste, and ξ_j is an iid econometric error term (Berry, 1994). Learning about microbead risks is expected to have a direct effect that shifts demand, and an indirect effect represented by a change in the dispersion of valuations among consumers of microbeaded toothpaste. We model this latter effect by recognizing that the boycott response term is a function of unobserved consumer heterogeneity through the distribution of preferences, F_s.

We capture demand-curve rotation by allowing each boycott-response parameter to be randomly distributed according to:

$$\delta_m = \delta_{m0} + \delta_{m1}\nu,\ \nu_m \sim N(0, 1), \tag{3}$$

where we interpret δ_{m0} as the direct effect of a boycott shock for non-beaded (m=1) and beaded (m=2) toothpaste, δ_{m1} is the indirect effect caused by changes in the dispersion of consumers' valuations for each toothpaste type, and ν_m is the estimate of variability in knowledge of the environmental risks associated with the use of microbeads. We allow the covariance matrix for the two random parameters to be non-zero.

Following the logic developed above, if $\delta_{20} < \delta_{10}$ then the demand for beaded toothpaste shifts leftward by more than the demand for non-beaded toothpaste during periods of consumer boycott. Similarly, if $\delta_{21} < \delta_{20}$ then the dispersion of consumers' valuations for beaded toothpaste decreases by more than for non-beaded toothpaste in response to the boycott, resulting in a counter-clockwise rotation in demand for beaded toothpaste relative to non-beaded toothpaste products. Such an outcome would imply that consumer demand for microbeaded toothpaste becomes more elastic relative to products not containing microbeads, a finding that would explain why retailers offered selective price discounts on products in the micro-beaded segment of the toothpaste market during the boycott period.

7.3 Structural Model Results

Table 5 presents our estimation results from the structural model of microbeaded toothpaste demand. The key parameters of interest are our estimated coefficients δ_{i0} and δ_{i1} that capture the shift in demand and the rotation effect, respectively, for toothpaste products. Based on our measure of Google search intensity as a proxy for information shocks on the microbead boycott, we interact the underlying variable in each case with a microbead-status indicator variable to capture the difference in the shift and rotation effects across these subcategories of toothpaste products.

Notice that demand decreases for beaded toothpaste and increases for non-beaded toothpaste during periods of high Google search intensity. Moreover, the decrease in demand for microbeaded toothpaste products is significantly larger than the overall decrease in demand for all toothpaste products, indicating that the consumer boycott was effective in reducing demand for toothpaste products containing microbeads.

The entries in Table 5 indicate that the dispersion of consumers' valuations for microbeaded toothpaste products declines significantly relative non-beaded toothpaste products during periods of high Google search intensity: $\delta_{21} < \delta_{11}$. The decrease in dispersion of consumers' valuations for microbeaded toothpaste products suggests a counter-clockwise demand rotation occurred for microbeaded toothpaste relative to non-beaded toothpaste products during periods of high Google search. This finding is consistent with our observation that retailers selectively discounted prices on microbeaded toothpaste products in response to the consumer boycott, which suggests that the boycott backfire was driven by retailer mediation through price changes as the elasticity of demand for beaded toothpaste products increased relative to non-beaded toothpaste products.

8 Conclusion

In this paper we have examined the effect of a consumer boycott on microbeads on equilibrium outcomes in the toothpaste category. We rely on variation in Google search intensity across the ten states that proposed bills to ban microbeads to identify the extent of consumer awareness regarding microbead pollution. We find that consumer demand for microbeaded toothpaste products decreases during periods of peak Google search intensity, but that retailers responded to the negative demand shock by selectively reducing prices of microbeaded toothpaste products and promoting them more frequently. As a result, the total quantity of microbeaded toothpaste products sold by retailers increased during the four-week period following the boycott, resulting in a boycott backfire.

A large literature on retail responses to demand shocks has documented that retailers reduce prices and promote products more frequently during periods of increased demand (Warner and Barsky, 1995; MacDonald, 2000; Chevalier et al., 2003). This finding suggests that retailers would raise prices for microbeaded toothpaste products and promote them less frequently in response to a negative demand shock. Our findings suggest the opposite outcome occurred. Retailers selectively discounted microbeaded toothpaste prices and promoted these product lines more frequently in response to the consumer boycott. The reason is that the boycott appears to have rotated the demand functions for beaded and non-beaded products, making demand for microbeaded toothpaste products relatively price elastic.

We confirm the intuition for this finding by examining changes in the dispersion of consumers' valuations during periods of peak Google search. During periods of peak search intensity, demand for all toothpaste products declines, while the dispersion of consumers' valuations decreases for microbeaded toothpaste products relative to non-beaded products. Our findings indicate that demand for microbeaded toothpaste products decreased, but became more price elastic relative to non-beaded products following the boycott. Retailers responded by offering selective price discounts on microbeaded toothpaste products, generating greater sales velocity for beaded products in the face of the consumer boycott.

An interesting area for future research is to examine the effect of retailer mediation of consumer boycotts in other contexts. An unusual aspect of the toothpaste market is that category-leading toothpaste manufacturers (Crest and Colgate) offered both microbeaded and non-beaded toothpaste products. This may have dampened the effectiveness of a consumer boycott relative to a boycott levied against an entire brand. While our findings provide evidence that a boycott backfire can occur when retailers mediate sales of boycotted products, further research is needed to examine a broader range of product categories and retail environments.

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Table 1: Summary Statistics

	median	mean	sd	min	max	N
Boycott Regression						
$\Delta Share14-13$	-0.000	-0.000	0.059	-1.220	1.228	1,221,689
$\Delta NumVariants14-13$	-0.000	-0.000	0.031	-0.481	0.485	1,221,689
$\Delta Price14-13$	0.001	0.001	0.354	-17.490	19.473	1,221,689
$\Delta Share14-13_{t-1}$	-0.000	-0.000	0.059	-1.220	1.228	1,221,689
County Environmental Score	77	64.722	29.172	0	98	1,221,689
First Week	0	0.250	0.433	0	1	1,221,689
High Google	0	0.084	0.277	0	1	1,221,689
High Google * Bead	0	0.018	0.133	0	1	1,221,689
Bead	0	0.216	0.411	0	1	1,221,689
Pricing Regression						
Price per Ounce	0.650	0.785	0.527	0.001	5.088	232,835
On Sale	1	0.505	0.500	0	1	232,835
First Week	0	0.247	0.431	0	1	232,835
High Google	0	0.100	0.300	0	1	232,853
High Google * Bead	0	0.016	0.125	0	1	232,835
Post	0	0.115	0.319	0	1	232,853
Post * Bead	0	0.018	0.132	0	1	232,835
Bead	0	0.155	0.362	0	1	232,835
Backfire Regression						
Ounces sold, log	5.534	5.556	1.481	0.000	11.589	429,855
First Week	0	0.249	0.432	0	1	429,855
High Google	0	0.088	0.283	0	1	429,855
High Google * Bead	0	0.043	0.203	0	1	429,855
Post	0	0.123	0.328	0	1	429,855
Post * Bead	0	0.061	0.239	0	1	429,855
Bead	0	0.495	0.500	0	1	429,855
Median Household Income	56,645	62,430	15,620	27,882	107,250	429,855

Table 2: Evidence of Boycott

	(1) All	(2) Low Green	(3) High Green	(4) DIDID	(5) With Trends
High Google	0.000856***	-0.0000600	0.00162***	0.000481	0.00109***
	(0.000192)	(0.000292)	(0.000253)	(0.00230)	(0.000203)
High Google * Bead	-0.00126**	0.000435	-0.00272***	-0.00320***	-0.00121**
	(0.000412)	(0.000627)	(0.000546)	(0.000549)	(0.000438)
$\Delta NumVariants14-13$	0.887***	0.912***	0.853***	0.887***	0.887***
	(0.00148)	(0.00199)	(0.00223)	(0.00148)	(0.00148)
$\Delta Price14-13$	-0.0334***	-0.0337***	-0.0332***	-0.0334***	-0.0334***
	(0.000130)	(0.000204)	(0.000168)	(0.000130)	(0.000130)
$\Delta Share14-13_{t-1}$	-0.0901***	-0.0833***	-0.0971***	-0.0900***	-0.0901***
	(0.000787)	(0.00112)	(0.00111)	(0.000787)	(0.000787)
First Week	0.0000547	-0.0000733	0.000164	0.0000504	0.0000362
	(0.000107)	(0.000161)	(0.000143)	(0.000107)	(0.000107)
Constant	-0.000219***	-0.000145	-0.000285***	-0.000218***	0.000107
	(0.0000557)	(0.0000833)	(0.0000749)	(0.0000557)	(0.0000990)
Brand-Event Interactions	No	No	No	Yes	No
Trends by Bead	No	No	No	No	Yes
Observations	1221689	565692	655997	1221689	1221689
R^2	0.287	0.323	0.249	0.287	0.287

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Notes: The table presents results from three fixed effects OLS regressions where a unit of observation is a brand-store-week from 10 states. Standard errors are in parentheses and are clustered by store. The panel starts the first week of January 2014 and ends four weeks after the first time the state's google search intensity passes 25. Columns two and three stratify the sample based on a county's lifetime environmental score. Low Green are counties with a lifetime score less than 75 and High Green are greater than 75. Column four uses a triple-difference specification, interacting the event by brand. Column five includes trends by Bead.

Table 3: Retailer Pricing Response to Boycott

	(1) Price per Oz	(2) Price per Oz	(3) On Sale	(4) On Sale
High Google	-0.00400	-0.00160	-0.00679	-0.0206
	(0.0135)	(0.0137)	(0.0272)	(0.0272)
High Google * Bead	0.0117**	0.00997	-0.0158	-0.00738
	(0.00405)	(0.00551)	(0.0135)	(0.0164)
Post	-0.0111	-0.0126	0.0129	0.0219
	(0.0104)	(0.0102)	(0.0226)	(0.0216)
Post * Bead	-0.00852*	-0.0101*	0.0404**	0.0483**
	(0.00330)	(0.00420)	(0.0151)	(0.0181)
First Week	0.00345	0.00153	-0.0613*	0.00444
	(0.0104)	(0.000778)	(0.0258)	(0.00368)
Constant	0.783***	0.779***	0.539***	0.506***
	(0.00894)	(0.00308)	(0.0218)	(0.00488)
Week of year	Yes	No	Yes	No
Brand-Event interactions	Yes	Yes	Yes	Yes
Trends by Bead	No	Yes	No	Yes
Observations	232835	232835	232835	232835
R^2	0.004	0.002	0.003	0.001

* $p < .05$, ** $p < .01$, *** $p < .001$

Notes: The table presents results from two fixed effects OLS regressions where a unit of observation is a UPC-chain-week in 10 states. Standard errors are in parentheses and are clustered by retailer-state. The panel starts the first week of January 2014 and ends four weeks after the first time the state's google search intensity passes 25. The price per ounce is in US dollars. The On Sale outcome is a dummy variable equal to one if that week's price is below the mode price, where the mode price is the mode over a six week moving window constructed similar to Chahrour (2011).

Table 4: Evidence that Boycott Backfires

	(1) All	(2) Low	(3) Middle	(4) High	(5) Trends
High Google	0.0297***	0.0171	0.0423***	0.0202*	0.0131***
	(0.00432)	(0.00892)	(0.00617)	(0.00792)	(0.00367)
High Google * Bead	-0.0572***	-0.0629***	-0.0608***	-0.0434***	-0.0250***
	(0.00436)	(0.00999)	(0.00613)	(0.00724)	(0.00471)
Post	-0.00254	0.00150	0.00105	-0.0178*	-0.0120***
	(0.00450)	(0.0102)	(0.00618)	(0.00863)	(0.00335)
Post * Bead	0.0317***	0.0284**	0.0277***	0.0427***	0.0663***
	(0.00364)	(0.0101)	(0.00497)	(0.00588)	(0.00413)
First Week	0.0547***	0.0580	0.0535***	0.0178	0.0291***
	(0.00995)	(0.0539)	(0.0141)	(0.0226)	(0.00147)
Constant	5.530***	5.061***	5.553***	5.870***	5.563***
	(0.00852)	(0.0533)	(0.0118)	(0.0207)	(0.00324)
Week of Year	Yes	Yes	Yes	Yes	No
Trends by Bead	No	No	No	No	Yes
Observations	429855	83138	236746	109971	429855
R^2	0.009	0.015	0.010	0.011	0.003

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Notes: The table presents results from four fixed effects OLS regressions where a unit of observation is a retailer-zip-bead-week from 10 states. The outcome variable is the natural log of the total ounces sold plus one. Standard errors are in parentheses and are clustered by store. The panel starts the first week of January 2014 and ends four weeks after the first time the state's google search intensity passes 25. The three columns on the right are subsamples based on the county's median household income level. Low income is defined as less than $50,000, about the 25th percentile, and High incomes as more than $75,000, about the 75th percentile.

Table 5: Random Coefficients Model Estimates

Variable	Estimate	Standard Error
Constant	0.638***	(0.0877)
Age	-0.000590	(0.00124)
Income	0.00112***	(0.000209)
Household Size	-0.0220***	(0.00594)
Hispanic	-0.0170	(0.0133)
Married	0.0165	(0.0135)
Education	0.00685**	(0.00230)
Employed	-0.0155	(0.0139)
Deal	-0.202***	(0.0302)
Bead	0.141^{+}	(0.0850)
Random Parameter Means		
High Google (δ_{10})	0.107*	(0.0511)
High Google * Bead (δ_{20})	-0.159*	(0.0634)
Random Parameter Std. Deviations		
High Google (δ_{11})	0.375***	(0.0395)
High Google * Bead (δ_{21})	0.192***	(0.0585)
Variance	-9.354***	(1.952)
Covariance	0.467***	(0.0422)
Month FE	Yes	
State FE	Yes	
Log lik.	-14426.2	
N	21825	

Standard errors in parentheses

* $p < .05$, ** $p < .01$, *** $p < .001$

Notes: The table presents results from a linear regression with random coefficients where a unit of observation is a household-brand-week from 10 states. The dependent variable is the price per ounce teh household paid, which proxies for their willingness-to-pay. Standard errors are in parentheses and are clustered by brand. The covariance between the random parameters is unstructured and the random parameters are estimated across brands. The panel starts the first week of January 2014 and ends four weeks after the first time the state's google search intensity passes 25.

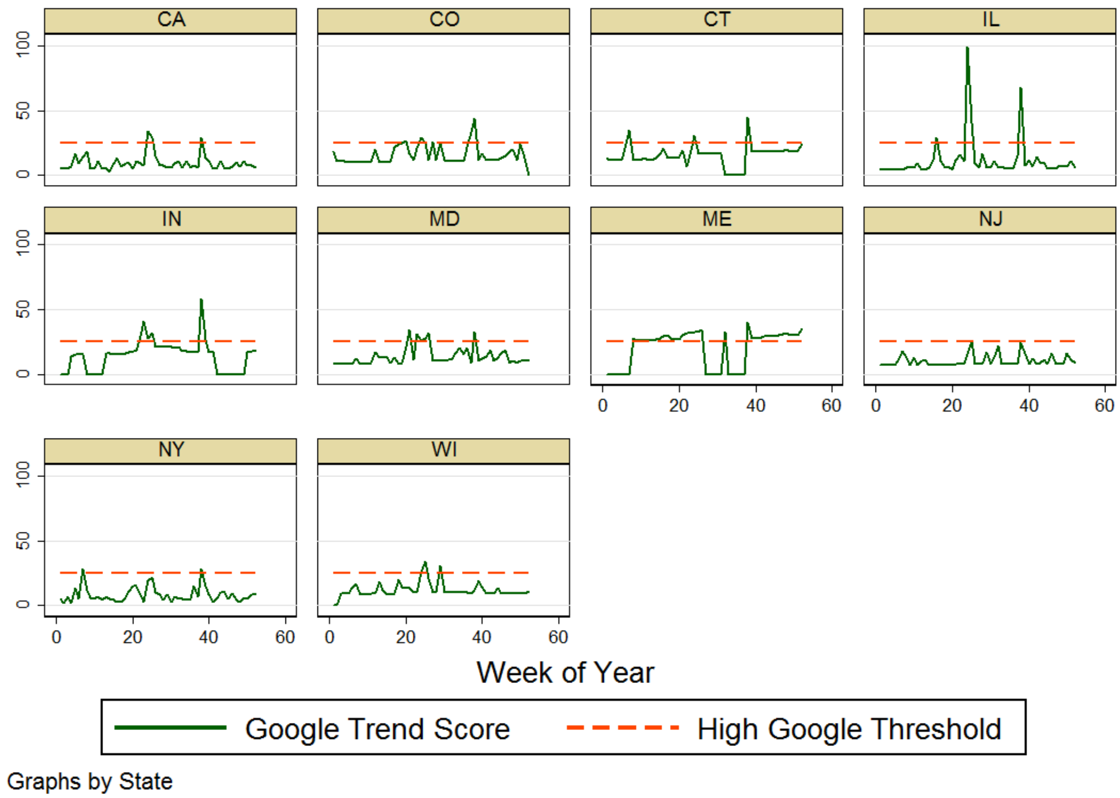

Figure I: Google Search Intensity by State

Notes: The figure presents the weekly google trends score for the search term "microbeads" for each week of 2014 for 10 states that had proposed legislation banning microbeads. The values are excerpted from a panel of google trends from Jan 6, 2013 to February 6, 2016. Each panel is the weekly score for that state. The threshold used to define a "High Google" event is a score of 25, shown as the dashed line in each panel.

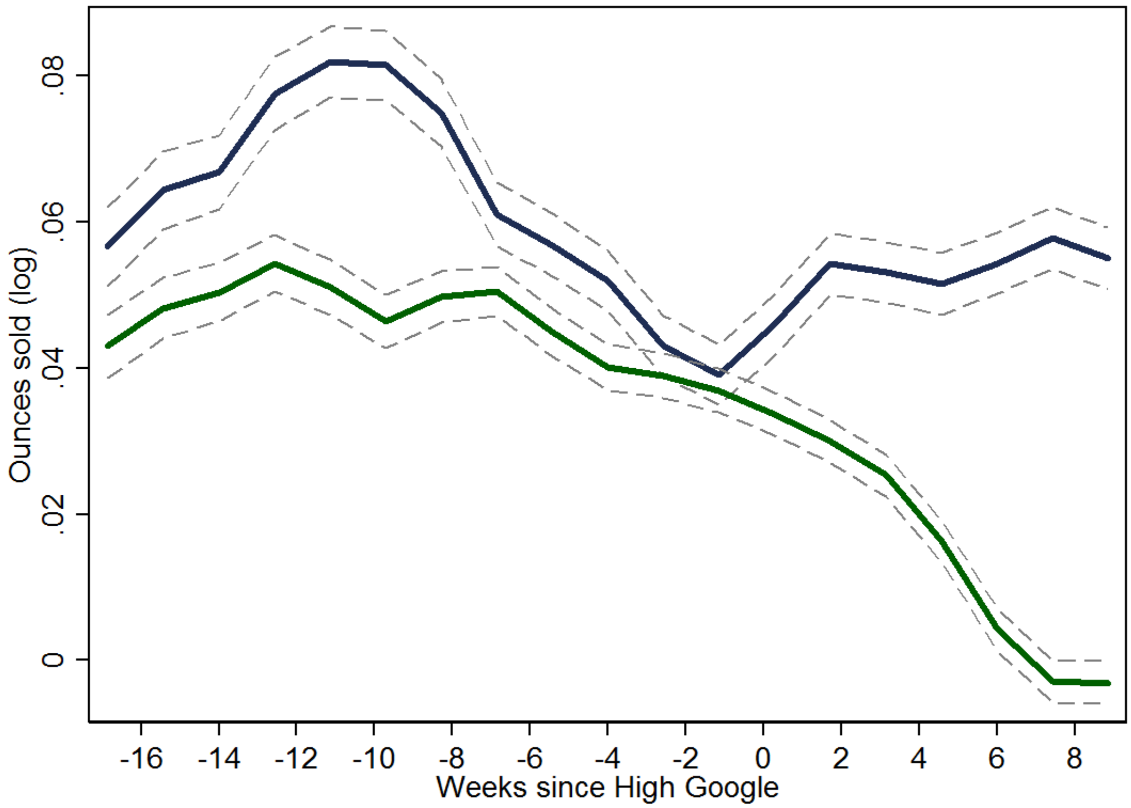

Figure II: Event Study Figure

Notes: The figure presents coefficient estimates from two local polynomial regressions where the outcome variable is the de-meaned natural log of ounces sold and data are aggregated to the store-bead-week level, as in Table 4. The green line is from a regression restricted to products without beads. The blue line is for products with beads. 95% confidence intervals are in grey. Time zero is the first week with a high google search value

A For Online Publication - Appendix

Table A1: Ingredients searched in Household Products Database

Ingredient Name
Nylon-12
Nylon-6
Polybutylene terephthalate
Polyethylene terephthalate
Polyethylene terephthalate (PET)
Polymethyl methacrylate
Polyethylene
Polypropylene
Polystyrene
Polytetrafluoroethylene (PTFE)
Polyurethane
Acrylate copolymer
Ethylene copolymer
Methacrylate copolymer
Styrene copolymer

Table A2: Brands and flavors with microbeads sold in Nielsen stores

Ingredient Name	Flavors
COLGATE LUMINOUS	ALL
COLGATE MAX CLEAN SMART FOAM	ALL
COLGATE TOTAL ZX PRO-SHIELD	ALL
COLGATE	ICY BLAST
CREST 3D WHITE	ALL
CREST 3D WHITE LUXE	ALL
CREST COMPLETE	FRESH MINT BLAST
CREST COMPLETE	EFFERVESCENT MINT
CREST COMPLETE	MINT
CREST PRO-HEALTH	ALL
CREST PRO-HEALTH FOR ME	ALL
CREST PRO-HEALTH FOR LIFE	ALL
CREST SENSITIVITY	ALL
CREST SENSI-REPAIR & PREVENT	ALL
CREST SENSI-RELIEF	ALL
CREST WHITENING EXPRESSIONS	EXTREME HERBAL MINT

Table A3: Beaded Toothpaste in Neighboring States

	Market Share Growth	
High Google	-0.0101***	(0.000904)
$\Delta NumVariants14-13$	0.898***	(0.0105)
$\Delta Price14-13$	-0.0361***	(0.00104)
$\Delta Share14-13_{t-1}$	-0.0166**	(0.00608)
First Week	-0.00410***	(0.000569)
Constant	0.000799***	(0.000147)
Observations	30615	
R^2	0.471	

Standard errors in parentheses

Obs are product line-store-week toothpaste in IL and WI.

Sample restriction: units sold every week of 2013-2014.

Panel starts Jan 2014 ends 4 weeks after IL first surge.

OLS fixed effects regression. SE clustered by store.

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table A4: Household Panel Summary Statistics

Variable	Mean	Std. Dev.
Price	0.623	0.489
Age	52.809	10.082
Income	70.934	36.354
Household Size	2.639	1.352
Hispanic	0.094	0.291
Married	0.700	0.458
Education	13.709	2.019
Employed	0.418	0.352
Deal	0.401	0.49
Bead	0.157	0.364
High Google	0.057	0.231
High Google * Bead	0.009	0.097
N	21825	

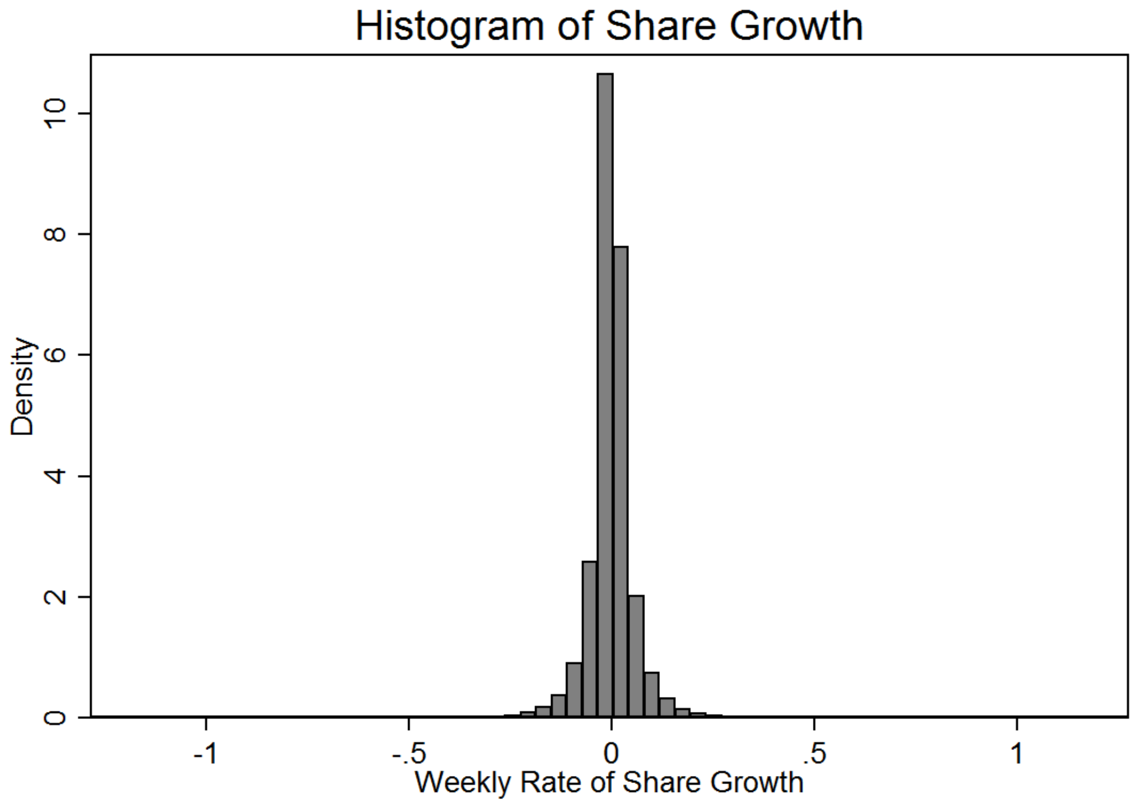

Figure AI: Histogram of Share Growth

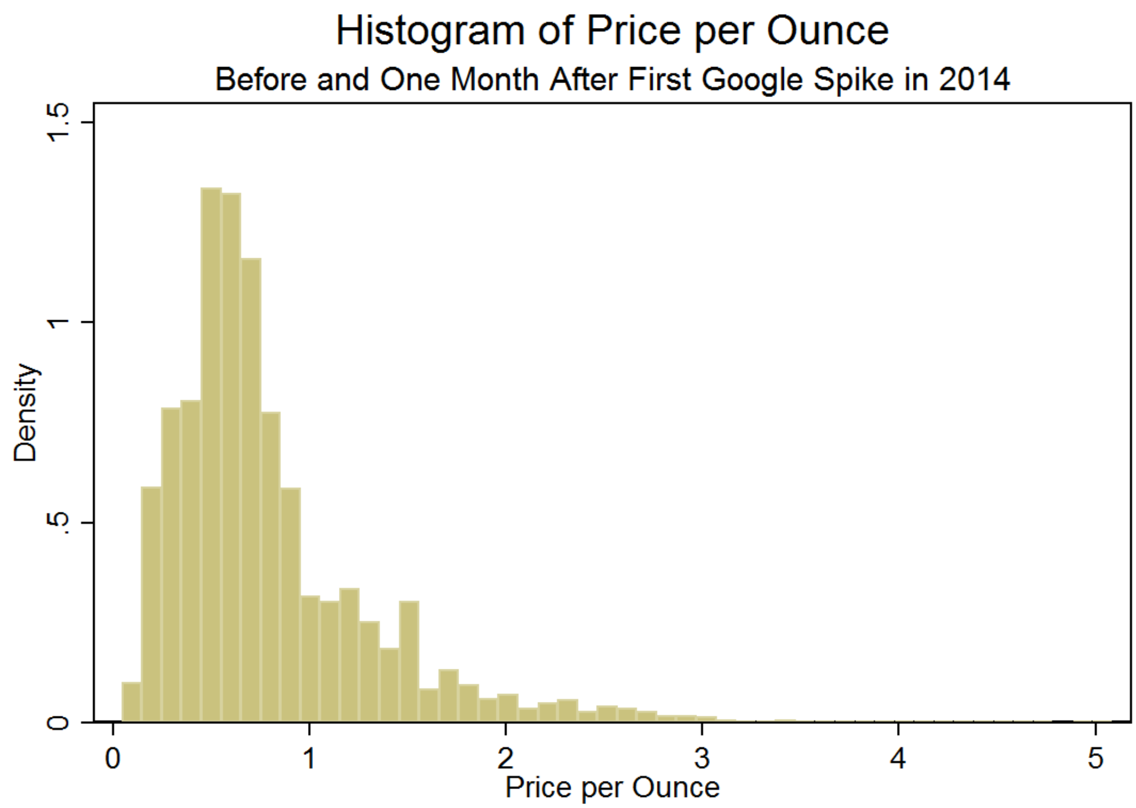

Figure AII: Histogram of Price Per Ounce

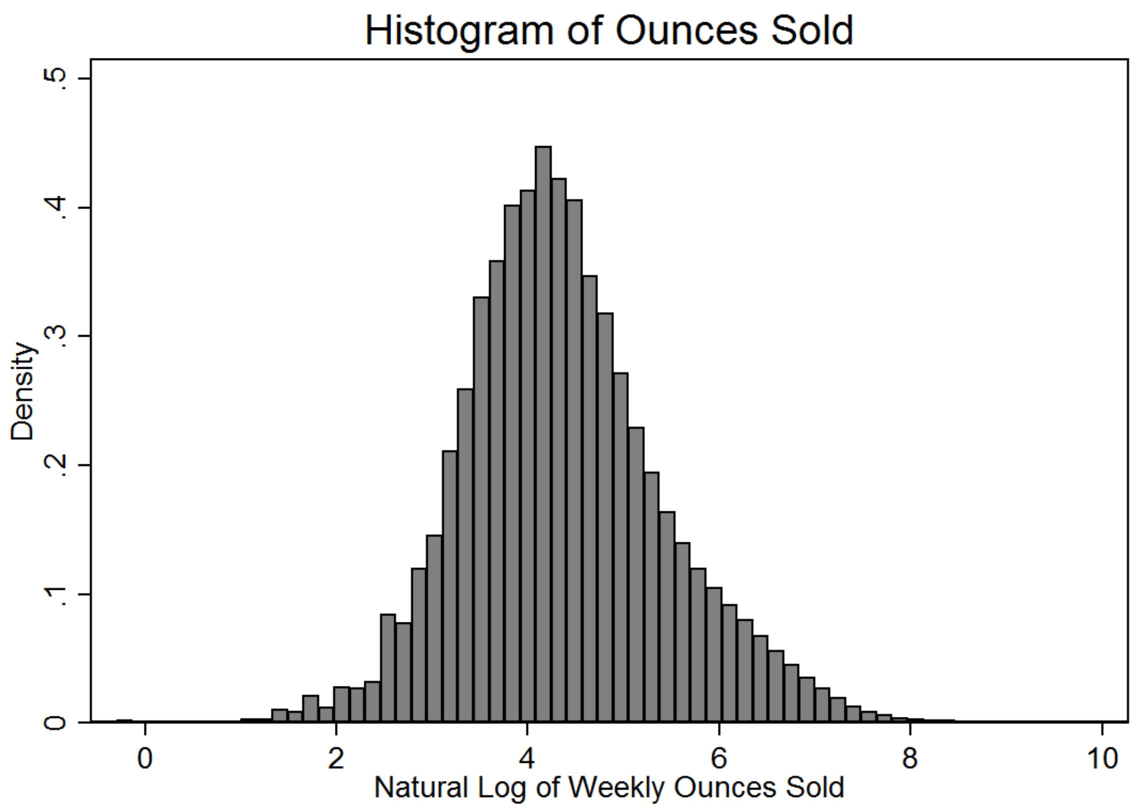

Figure AIII: Histogram of Ounces Sold