Why Does Popcorn Cost So Much At the Movies?
An Empirical Analysis of Metering Price Discrimination[1]

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Abstract

Prices for goods such as blades for razors, ink for printers and concessions at movies are often set well above cost. This paper empirically analyzes concession sales data from a chain of Spanish theaters to demonstrate that high prices on concessions reflect a profitable price discrimination strategy often referred to as "metering price discrimination." Concessions are found to be purchased in greater amounts by customers that place greater value on attending the theater. In other words, the intensity of demand for admission is "metered" by concession sales. This implies that while some consumers' surplus may be reduced by the high concession prices, other consumers on the margin of attending may benefit from theaters' decisions to shift their margins away from movies and toward concessions.

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1. Introduction

When a variable unit good is sold after the purchase of a single unit good, the price of the variable unit good is often observed to be well above cost. For instance, popcorn purchased after entering a movie theater, sports stadium or other venue charging admission is priced much higher than in grocery stores, small shops or restaurants. A common presumption is that the venues exploit the fact that customers have little if any choice between concession sellers. While this presumption is probably accurate, it is important to recognize that high concession prices might reflect the decision of the venue to shift profits from admission tickets to concessions. This strategy has been termed metering price discrimination because the surplus extracted from a customer is "metered" by how much of the aftermarket good they demand.

Like other forms of price discrimination, metering has the ability to increase efficiency because it can open access of a good to customers that would otherwise be priced out of the market. For example, if a venue priced concessions at or near marginal cost, it would extract all of its value in the admission price, thereby leaving some customers out. So, while the surplus of some consumer may be reduced by high concession prices, total surplus, producer surplus and the surplus of other consumers may be increased.[2]

For high aftermarket prices to be associated with efficiency increases, the primary good price must be predicted to be lower than it would be under a competitive aftermarket. However, primary good prices are only lower because of metering if customers that demand more aftermarket goods (e.g., concessions) also place a greater value on the primary good (e.g., admission). This necessary demand condition has been shown by Oi (1971), further explored by Schmalensee (1981), and applied to the case of admission tickets and concessions by Rosen and Rosenfield (1997). Furthermore, the explanation of

[2] One condition under which total surplus could decrease, even though movie admission increases, is if the reduced concession sales from high concession prices reduces surplus more than the surplus increase for admission. We doubt this is possible in this case and many other similar examples because the aftermarket goods (concessions) are of substantively less consequence than the primary good (movie admission). Specifically, concessions could be purchased without ever attending a movie, so it seems reasonable to weight the surplus effects of the movies much greater than those of the concessions.

metering has been applied to many goods such as razors and blades and Polaroid cameras and film, and has been a common efficiency rationale for the decision to tie aftermarket goods to the purchase of primary goods (see Peltzman, 2005 and Klein, 1996).

Despite the awareness of metering and its necessary demand conditions, there has yet to be any work estimating demand to test whether it can in fact explain high aftermarket prices. This paper fills this void by analyzing a unique data set with approximately five years of weekly attendance, box office revenue and concession revenue for a chain of 43 Spanish movie theaters. Our data is aggregated at the week and theater level, but we define an estimation strategy that allows us to uncover the joint distribution of the individuals' willingness to pay for movie admission and demand for concessions. Our identification strategy appeals to the intuitive notion that attendees in lower attendance weeks may enjoy going to movies more than attendees in average weeks. In other words, we define conditions under which low attendance proxies for customer types with greater values for admission. In such cases, a negative correlation between attendance and concession demand per person in the data results from a positive relationship between willingness to pay for admission and concession demand.

We therefore estimate this demand relationship to evaluate the theater's pricing incentives, while controlling for a number of other factors that could influence the correlation between concession demand per attendee and attendance. Perhaps the most significant confounding factor is the fact that concession lines might be longer in peak demand periods. This could naturally create a negative correlation between attendance and concession sales per attendee. We address this concern in a few ways. First, we estimate this correlation for the set of weeks in which attendance is less than the average week. This excludes the peak weeks when lines are most likely to be long. Second, we include a variable that proxies for the length of concession lines: the theater's forecast error for attendance. A unique feature of the data set is that we observe the week-ahead demand forecasts that the theater uses to set staffing for concession lines. By measuring how actual realized demand differs from the forecasted demand, we can proxy for the length of concession lines. Third, we allow the relationship to differ by every decile of attendance at the theater. This is because staffing concerns can handle some degree of queuing, but theaters also have fixed inputs (e.g. soda machines and registers) that can lead to lines in very busy times even when staffed appropriately. Together, we believe these factors account for any queuing that could confound the relationship between concession sales per attendee and attendance.

The empirical analysis confirms that there is a negative correlation between concession demand per attendee and attendance. The finding is robust to the concession line proxies described above, as well as fixed effects for theaters and weeks. In specifications where we allow the relationship to differ by decile of attendance at each theater, we find that concessions sales per person is even lower in the top decile weeks when the fixed inputs problem may create long concession lines. Furthermore, this finding suggests that any confounding factor must persist across all attendance levels rather than reflecting systematic unobserved differences between high and low attendance movies.

For a limited sample of the data, we also have movie characteristics to control for horizontally differentiated preferences across genres, ratings classes, US box office performance of movies and the number of weeks since release. Through all specifications with theater fixed effects, we find that concession sales per person are negatively correlated with attendance.

In addition to this, we find that higher average ticket prices paid (reflecting more regular admissions than senior and student discounts) are associated with greater concession sales. This implies that even in the pricing dimension through which a theater can identify high and low willingness to pay customers, we still find that higher willingness to pay customers (those paying higher ticket prices) consume more concessions on average.

To summarize, we consistently find that willingness to pay for movies is positively correlated with concessions demand. The finding is supported by a negative relationship between attendance and concessions sales per attendee because high attendance weeks tend to also attract customers with lower average valuations for the theater and these weeks are associated with lower concession sales per person. Further support for the positive relationship between willingness to pay and concession demand comes from the finding that when more students and seniors arrive at the theater (and pay lower prices due to their presumed lower willingness to pay for admission), we find that concession sales are lower. This documented demand relationship therefore suggests that theaters should be shifting their margins from ticket prices to concessions and that if they are doing this, the high aftermarket prices may be allowing them to admit more people to their movies.

The rest of the paper is organized as follows. The next section describes the motivation behind our empirical analysis. Section 3 describes the data. Section 4 discusses our empirical approach and results and section 5 concludes.

2. Motivation for Empirical Analysis

The existing theoretical work (e.g. Oi, 1971; Littlechild, 1975; Schmalensee, 1981; and Rosen and Rosenfield, 1997) is instructive about the joint distribution of demands necessary for sales of an aftermarket good to profitably meter the variation in willingness to pay for the associated primary good. However, these theoretical models do not provide intuition about how to uncover this joint distribution from available data. In this section, we illustrate how variation in a single vertical attribute specific to the primary good can uncover the correlation between willingness to pay for the primary good and demand for the aftermarket good.

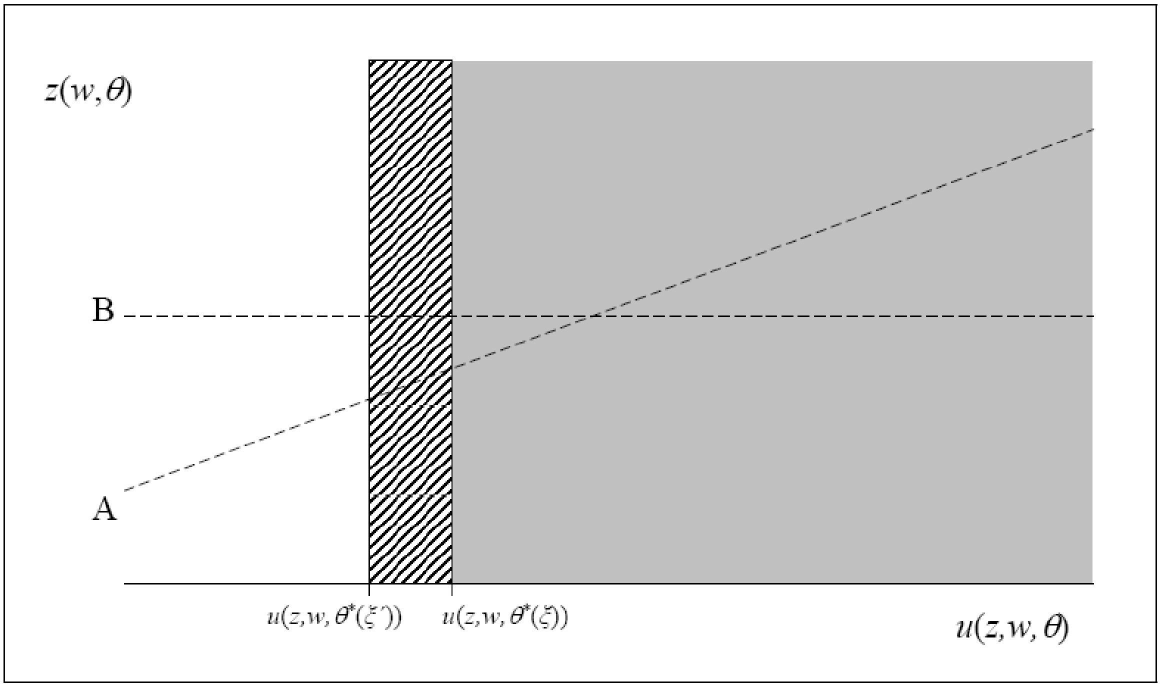
We begin by defining the indirect utility obtained when purchasing the primary good to be $v_1(z,w,\theta,\xi_1)$, while the indirect utility when not purchasing the primary good is $v_0(\theta,\xi_0)$. z is concessions consumed by an individual, w is the price concessions, θ is a parameter (or vector of parameters) representing tastes, and ξ_1 and ξ_0 are variable attributes common to all individuals that respectively affect utility when purchasing and not purchasing the primary good. We assume the following properties of the indirect utilities:

$$\partial v_1 / \partial w < 0, \quad \partial v_1 / \partial \xi_1 > 0, \quad \partial v_0 / \partial \xi_0 > 0 \tag{1}$$

Define $v(z,w,\theta,\xi) = v_1(z,w,\theta,\xi_1) - v_0(\theta,\xi_0)$ to be the surplus when purchasing the primary good. The demand function for concessions is defined to be $z(w,\theta)$. Note that we have assumed that the variable vertical attributes do not enter $z(w,\theta)$. We discuss below why this assumption is necessary for identification. We also define $u(z,w,\theta)$ to be the willingness to pay, net of the variable attributes, ξ. An obvious example would be $v(z,w,\theta,\xi) = u(z,w,\theta) + \xi_1 - \xi_0$. $u(z,w,\theta)$ represents a measure of willingness to pay for the primary good that is invariant to changes in ξ. Variation in $u(z,w,\theta)$ across individuals in the population is therefore indicative of heterogeneity in the willingness to pay for the primary good, where $u(z,w,\theta)$ indexes this heterogeneity. In our identification below, we focus on variation in ξ's that do not disrupt the property that the ordering of $u(z,w,\theta)$'s characterize the ordering of $v(z,w,\theta,\xi)$'s. An example of such a ξ would be price. Price variation affects the surplus, $v(z,w,\theta,\xi)$, but does not affect $u(z,w,\theta)$, and, we expect that price variation would never lead a lower type customer, $u(z,w,\theta_L)$, to have a greater surplus than a higher type customer, $u(z,w,\theta_H)$. ($v(z,w,\theta_L,\xi) < v(z,w,\theta_H,\xi)$ for all ξ and all $\theta_L < \theta_H$.)

The existing theory implies that a positive correlation between $v(z,w,\theta,\xi)$ and $z(w,\theta)$ is necessary for metering. A positive correlation between $u(z,w,\theta)$ and $z(w,\theta)$ also satisfies the condition. To clarify how variation in ξ's can identify this correlation, refer to Figure 1, which depicts the joint distribution of $u(z,w,\theta)$ and $z(w,\theta)$. The conditions for profitable metering are satisfied if the dashed line labeled A in Figure 1 represents the relationship between $u(z,w,\theta)$ and $z(w,\theta)$. Theory shows that metering would not be profitable if the relationship were represented by the dashed line labeled B (i.e. a zero correlation between aftermarket demand and willingness to pay for the primary good). In such a case, a firm would be better off pricing the aftermarket good at its constant marginal cost to create as much surplus as possible, while extracting surplus through the price of the primary good.

Figure 1



We define the marginal buyer of the primary good to be $\theta^*(\xi)$ such that $v_1(z,w,\theta,\xi_1) = v_0(\theta,\xi_0)$ and $u(z,w,\theta^*(\xi))$ is the marginal consumer's willingness to pay for the primary good, net of ξ. Identification of the joint distribution of $u(z,w,\theta)$ and $z(w,\theta)$ can be obtained if we have variation in the common vertical attribute, ξ. The most obvious source of variation to use is exogenous price variation. For instance, if ξ_1 is price and we observe a reduction in price from ξ_1 to ξ_1', the marginal consumer, defined by $u(z,w,\theta^*(\xi))$ changes to $u(z,w,\theta^*(\xi'))$. This implies a couple things. First, the price reduction increases attendance to include both the gray shaded area and the region with the diagonal markings. Second, the average willingness to pay for the primary good across buyers, $E[u(z,w,\theta) \mid u(z,w,\theta) > u(z,w,\theta^*(\xi))]$ decreases. Together these tell us that we can infer variation in the willingness to pay from variation in demand for the primary good: observations with greater than average primary good demand due to ξ imply buyers with a lower than average willingness to pay. To evaluate whether willingness to pay for the primary good is positively correlated with demand for the aftermarket good, we then only need to observe whether aftermarket demand per buyer, $E[z(w,\theta)|u(z,w,\theta) > u(z,w,\theta^*(\xi))]$, decreases when primary good demand increases. In Figure 1, primary good demand increases are captured in the growth of the shaded area, while aftermarket demand decreases are captured by the shrinking ratio of the area under line A divided by the shaded area.[3]

In data, we only need to observe the demand for aftermarket and primary goods, Q and Z. The ratio, Z/Q , reveals $E[z(w,\theta)|\ u(z,w,\theta) > u(z,w,\theta^*(\xi))]$. While we cannot readily observe $E[u(z,w,\theta) \mid u(z,w,\theta) > u(z,w,\theta^*(\xi))]$, the argument above implies that increases in Q correspond to decreases $E[u(z,w,\theta) \mid u(z,w,\theta) > u(z,w,\theta^*(\xi))]$. The correlation can then be revealed through the correlation of Z/Q and Q.

While finding exogenous price variation is always challenging, the fact that prices are fixed over time for our empirical application of movies (Orbach and Einav, 2007) eliminates price from the set of potential factors that can reveal variation in willingness to pay. On the other hand, while prices are fixed, variation in other vertical attributes, holding all else equal, can reveal willingness to pay. In our empirical analysis, we appeal to two types of vertical attributes: (i) those affecting the utility when attending a movie, ξ_1; and (ii) those affecting the outside opportunity of not attending a movie, ξ_0.

An example of an attribute that affects utility when attending a movie is the variation in quality, conditional on other factors that capture horizontally differentiated preferences for a movie. Referring back to Figure 1, an increase in movie quality, holding all else constant, would also reduce $u(z,w,\theta^*(\xi))$ to $u(z,w,\theta^*(\xi'))$ and increase attendance, allowing the joint distribution of valuations and concessions to be revealed from the correlation of Z/Q and Q, so long as the quality shock satisfies two conditions: a) it does not systematically change the ordering of valuations along the horizontal axis; and b) it does not affect demand for concessions (i.e. $z(w,\theta)$ does not include ξ). In our empirical analysis below, we use week and year fixed effects as well as characteristics of movies playing in a theater in a given week to capture horizontal variation in preferences for movies, such that variations in attendance can be interpreted as variation in vertical quality that satisfies conditions a) and b). We should note that our approach does not rule out additional sources of horizontal preference variation, we only require that such

[3] When ξ is interpreted as price, as in this example, the fact that ξ does not enter $z(w,\theta)$ assumes away income effects. This assumption is common in the theoretical literature, though Schmalensee (1981) explores the sensitivity of the findings to the inclusion of income effects.

horizontal preference variation is not systematically related to either high or low attendance. Specifically, our empirical analysis below estimates the slope of the dashed line (e.g. A or B) in Figure 1 across 10 regions of attendance (depicted along the horizontal axis). If systematic horizontal demand shocks correlated with concessions explain the relationship we eventually estimate, these systematic differences between higher and lower attendance movies would have to recur within each of these ten regions.

While we have focused primarily on shocks that affect movie quality, ξ_1 , until now, shocks to outside opportunities could also allow attendance variation to reveal variation in customer types. Local festivities or other entertainment opportunities are examples. Another obvious example affecting outside opportunities is weather. Specifically, weather that decreases the utility of outside alternatives can be interpreted to decrease ξ_0 to ξ_0', such that $u(z,w,\theta^*(\xi))$ also reduces to $u(z,w,\theta^*(\xi'))$. While weather is an appealing shifter of $u(z,w,\theta^*(\xi))$, it is possible that weather might also affect concession demand (i.e. $z(w,\theta,\xi) \neq z(w,\theta)$) if, for instance, movie-goers consume more or less concessions in hot, cold or rainy weather. We therefore use weather in our empirical analysis below as an additional control to assure that observed variation in attendance is not driven by this factor which could be confounded with concession demand.

We can intuitively describe the identification as follows. If a theater has a poor set of movies that lowers its attendance below average, or if the outside alternative improves, then the marginal customer from the average week will no longer attend. If we also observe average concession sales per attendee to increase, it tells us that the customers opting not to attend in the week with below average attendance must have consumed fewer concessions per person than those individuals that still attend. Identification of this relationship would imply that firms should charge premiums on concessions rather than extracting all consumer value through admission prices.

One valuable aspect of our approach is that we can uncover the underlying taste distributions of consumers for movies and concessions without many of the parametric assumptions required in a structural approach to estimation. Most empirical demand analyses of price discrimination use a structural approach in which a utility function is specified as a function of parameters, then the population distribution of the parameters is estimated from the data (e.g. Leslie (2004), McManus (2000), Cohen (2000) and Hartmann and Viard (2006)). In our case, we motivate our empirical approach with a flexible utility function defined over the two goods. This approach allows us to obtain inference by estimating fixed effects regressions that uncover the necessary characteristics of the joint distribution of movie and concession demands without specifying the functional form of the utility function or the distribution of model parameters. In other words, our estimates hold for various utility functions.

Our non-structural approach is related to "reduced-form" empirical analyses of price discrimination, but is substantively different in emphasis. The disadvantage of our non-structural approach is that while we can validate the profitability of a metering price discrimination strategy at observed prices, we do not have estimates of model parameters that allow us to consider counterfactual pricing policies. This is common to other non-structural approaches that have been used to empirically analyze price discrimination (e.g. Shepard, 1991; Miravete and Röller, 2004; Seim and Viard, 2004; Busse and Rysman, 2005; Borzekowsi, Thomadsen and Taragin, 2006). However, our work differs from these papers in that most of these relate the incidence of price discrimination to market structure.

3. Empirical Application and Data Description

We evaluate whether metering price discrimination is a motivating factor for high margins in aftermarket goods within the context of concession sales at movie theaters. While there is a growing literature studying the economics of the movie industry,[4] Gil and Hartmann (2007) is the only other paper that analyzes actual concession sales data. That paper documents stylized facts and trends in concessions, but does not consider the economic incentives behind concession pricing.

The data we use consists of weekly concession sales, box office revenues and attendance from a Spanish exhibitor. The data spans from January 2002 to June 2006 and contains

[4] Empirical papers in this area have considered a wide array of topics such as the vertical structure of movie exhibition (Gil, 2004), the location of theaters (Davis, 2006), release decisions (Krider and Weinberg, 1998; Elberse and Eliashberg, 2003; Einav, 2006;), run-length decisions (Eliashberg et.al. 2001; Ainslie, Dreze and Zufryden, 2005), financing decisions (Goettler and Leslie, 2005) and risk and uncertainty (De Vany, 2004).

information on 43 different theaters during that time. These 43 theaters are in 30 different cities in 17 provinces.

Even though we observe 43 different theaters during the 5 years of data, we do not observe 43 theaters at all times since the Spanish exhibitor sold a few theaters, built up new theaters and acquired theaters from other exhibitors that exited the market. The sample starts with 24 theaters and ends with 37 theaters. The missing six theaters at the end of the sample were mainly old theaters located in Barcelona & Madrid downtown. Most of these theaters were not directly owned by the exhibitor, just operated. In these cases, the owners of the property decided to sell the locations for other uses (housing, supermarkets or even nightclubs). New theaters are made up of both newly constructed theaters and newly acquired theaters.[5]

Since we focus on the study of concession sales, we exclude from our analysis those theaters for which the concession sales are outsourced.[6] After dropping those theaters, we are left with 6,206 weekly observations from 43 different theaters. These theaters differ in size and seating capacity. The theaters in our sample have from 1 to 24 screens and range from 396 to 5,300 seats. Detailed summary statistics are available in Table 1A.

Table 1A also provides summary statistics for other variables used in our analysis. Weekly attendance varies from 348 to a bit over 40,000 attendees with an average close to 8,900. These numbers denote the skewness of the distribution of attendance across theaters. Table 1A also summarizes the forecast error and weekly weather for each theater. The forecast error is defined as the actual attendance minus the week-ahead forecast which is used to determine staffing of concession stands. Large forecast errors should therefore proxy for long concession lines. We observe the weather data for most of the observations, however our data source was missing data for many cities during the

[5] To put this in a historical perspective, the Spanish market was no different than other western economies in the late years of the 1990s and beginning of 2000s in that it experienced a rapid growth in number of theaters (and screens). This growth came both from new developments and new exhibitors coming in the market. After such rapid growth, the movie demand did not respond as industry managers had first anticipated and exhibitors were required to cut losses and investment. This manifested in closing of older theaters, cancellation of new projects and firm exit. The latter caused a major consolidation in the industry where surviving firms acquired a number of theaters operated until that moment by other exhibitors.

[6] As pointed out earlier, some of these theaters are owned by independent real estate companies. In some of these cases, the theater owners honor long-term contracts with outside concession providers. In these cases, concession sales are not available and therefore we drop these theaters from the study.

month of January 2004. Rain days within the week vary from 0 to 8, with the eight arising because the final week of one year is classified to have 8 days and rain was observed on all 8 days.

The data also shows that the average concession spending per attendee is close to 1.6 euros and ranges from 0.24 to almost 2.94 euros. Box Office per person averages 4.7 and ranges from 2.6 to 6.3 euros. This variable deserves further clarification since it provides information on what type of customer is entering the theater in any given week.

This firm follows a rather distinct pricing schedule. The firm charges three different prices throughout the week. We can call these different prices a high price p^H, a non-peak price p^L and a discount price p^S. The firm charges p^H to all individuals attending theaters on Saturday and Sunday (and festive days). On Wednesday, the theater charges the discounted price p^S to all attendants. Finally, the other days during the week (Monday, Tuesday, Thursday and Friday) the theater enforces third degree price discrimination. During these days, the theater charges the discounted price p^S to students and seniors, and the non-peak price p^L to all attendants that do not identify themselves as students or senior citizens.

Therefore variation in Box Office per person brings information on what type of individuals are attending the theater in any given week compared to other weeks. For example, an increase in Box Office pp (average ticket price) means that a higher share of attendants arrives during the weekend or that a lower share of students and senior citizens is attending the theater and therefore tells us information on the average willingness to pay of the individual attending the theater.

We also use screening data from Gil (2004). These data provide information on what movies each theater is playing for the first 26 weeks during the year 2002. During those weeks, we only have data for 24 theaters that differ in size from 2 to 16 screens and in seating capacity from 396 to 3875 seats. See Table 1B for detailed summary statistics. We use information on movie characteristics such as movie genre, rating classification, weeks after release and US box office revenue of the movie. To merge these into weekly theater observations we weight each movie's characteristics by its total Spanish box office revenue across all weeks. We see that theaters typically have more adventure movies and PG13 movies than other genres or classifications. We also see that the weighted average weeks after release are 6.21, the weighted average share of opening films is 0.13 and there are about 2 movies opening in a given theater week. US box office revenue is reported in millions and theaters weekly movie offerings have a weighted US box office revenue average of $185.75 million.

4. Empirical Methodology and Results

We now analyze the data to evaluate the efficacy of using concession sales to price discriminate across customers with different valuations for movies. The work of Rosen and Rosenfield (1997) and Schmalensee (1981) documented that if marginal attendees demand fewer less concessions then firms would have an incentive to price concessions above marginal cost. We therefore assess how concession sales per person vary as demand shocks lure or deter the marginal theater attendee. We use a variety of fixed effects or other explanatory variables to assure that this relationship is not driven by composition effects. Specifically, we want to be sure that movie-specific effects or other demand shocks are not altering the entire composition of attendees.

Empirical Methodology

In this section, we describe how traditional price discrimination in movie admission tickets (e.g. student and senior discounts as well as discount days or shows) both affects the identification intuition described in section 2 and provides an additional test for whether customers with a greater willingness to pay for admission also demand more concessions.

Our primary variable of interest is average concession revenue per attendee, $AR^{CO} = pZ/Q$. Given that this aggregates over the pricing classes, $j \in \{L, H, S\}$, described above, it is useful to decompose AR^{CO} as follows:

$$AR_{it}^{CO} = \frac{p^{CO}\left[Z^H\left(Q^H\right) + Z^S\left(Q^S\right) + Z^L\left(Q^L\right)\right]}{Q^H + Q^S + Q^L} \tag{2}$$

p^{CO} is the price of concessions. For simplicity, and due to data limitations, we will assume that there is a single uniform price for concessions. Recall from the data description above that there are three types of customers that enter a theater: Q^L is the demand from customers that do not have third-degree price discrimination discounts, but do elect to visit the theater in non-peak periods to pay lower ticket prices; Q^H is the demand from customers that elect to visit the theater in a peak-demand period such as a weekend and may or may not have access to third-degree discounts in other periods; Q^S is the demand from customers such as students or seniors that attend in periods when they can realize their discounts. $Z^j\left(Q^j\right)$ is the total concession demand from customers that paid price j, where the function allows this demand to be increasing or decreasing with the total number of attendees in price category j. By the arguments described in section 2, if $\partial \frac{Z^j\left(Q^j\right)}{Q^j} \Big/ \partial Q^j$ is less than zero, then the marginal customer of type j does consume fewer concessions and it will be profitable to charge a premium for concessions.

In our data, we do not observe the demand of each type of customer, Q^j, but we do have information about the relative size of each group as observed through the box office revenue per person, AR^{BO}, where:

$$\begin{aligned} AR_{it}^{BO} &= \frac{p^H q^H + p^S q^S + p^L q^L}{q^H + q^S + q^L} \\ &= p^H \alpha^H + p^S \alpha^S + p^L \left(1 - \alpha^H - \alpha^S\right) \end{aligned} \tag{3}$$

α^j is the share of attendees that paid price j. While we also do not directly observe the α^j s, AR^{BO} informs us whether there is a relatively larger or smaller fraction of customers that pay the full price, p^H. We now redefine our dependent variable in terms of the α^j s as well:

$$AR_{it}^{CO} = p^{CO}\left[AQ^{CO,H}\left(Q^H\right)\alpha^H + AQ^{CO,S}\left(Q^S\right)\alpha^S + AQ^{CO,L}\left(Q^L\right)\left(1-\alpha^H-\alpha^S\right)\right] \tag{4}$$

where $AQ^{CO,j}\left(Q^j\right) = {Z^j\left(Q^j\right)}/{Q^j}$. Under this specification, if ${\partial AQ^{CO,j}}/{\partial q^j}$ is less than zero, charging a premium on concessions to price discriminate will be profitable.

Using equation (4), we can more specifically define the relevant empirical relationships in the data. Our primary relationship of interest is the correlation between AR^{CO} and attendance. If $Corr\left(AR^{CO}, Q\right) = 0$, then we can infer that customers of each type consume a constant amount of concessions, Ω^j, such that $AQ^{CO,j}\left(Q^j\right) = \Omega^j Q^j$.[7] Under this null hypothesis, it would not be profitable to raise the price of concessions to extract more revenue from customers with higher movie values.

If this null hypothesis is rejected, a negative sign of this correlation will support the use of a premium on concessions to price discriminate, while a positive sign will suggest that the practice may not be appropriate for the purposes of price discrimination. Once again, while we cannot measure the sign of each ${\partial AQ^{CO,j}}/{\partial Q^j}$, we will evaluate the average effect. This could be rationalized by an assumption that the signs are identical for all types, but this assumption is not necessary for the average effect to indicate the profitability of the price discrimination practice. Clearly a very negative ${\partial AQ^{CO,j}}/{\partial Q^j}$ could offset a modestly positive ${\partial AQ^{CO,k}}/{\partial Q^k}$ to make the price discrimination profitable.

Finally, to simplify the specification and avoid an explanatory variable, Q, in the denominator of the dependent variable, $AR^{CO} = {R^{CO}}/{Q}$, we transform these variables using logs such that our specifications are of the form:

$$\ln(R^{CO}) = \beta_Q \ln(Q) + \beta_R AR^{BO} + \beta_x X + \varepsilon \qquad (5)$$

[7] One exception to this would be if increasing concession consumption for one type of customer were perfectly offset by decreasing concession consumption from another type of customer. This coincidence seems unlikely and could be ruled out by assuming that ${\partial AQ^{CO,j}}/{\partial q^j}$ had the same sign for all types j.

β_Q is therefore interpreted as the percentage increase in concession revenue resulting from a one percent increase in attendance. If $\beta_Q < 1$, we infer that concession revenue per person is decreasing with attendance and that theaters should in fact price concessions above marginal cost.

The presence of AR^{BO} in the above specification serves two purposes. First, it controls for differences in the composition of ticket prices paid to avoid confounding estimates of β_Q. Second, the coefficient β_R is itself indicative of whether customers with a greater willingness to pay, as identified by paying a higher ticket price, consume more concessions than those customers paying a lower ticket price.

Results

We now begin to analyze this relationship. The first column of Table 2 reports the simple regression of $\ln(R^{CO})$ on $\ln(Q)$ and AR^{BO}. In this and all other specifications, *'s indicate significance from zero for all variables except $\ln(Q)$, in which case *'s indicate significance from 1. The estimated coefficient on $\ln(Q)$ is significantly greater than one, but this is primarily due to systematic differences in theaters as is clear from specification (2) which controls for the number of screens and the number of seats per screen at each theater. To account for other unobservable theater characteristics, specification (3) includes theater fixed-effects and reveals a coefficient on $\ln(Q)$ that is significantly less than one. This is the primary result suggesting that theaters should charge premiums on concessions to meter willingness to pay for admission. Note that the coefficient on AR^{BO} is also positive suggesting that groups with identifiably greater willingness to pay for admission consume more concessions per person. The remainder of the specifications illustrates the robustness of these results to a variety of potentially confounding factors.

Specification (4) in Table 2 includes week fixed effects to account for seasonality factors such as annually recurring summer or holiday weeks. We see that the signs of the coefficients of interest are unchanged and the effects become stronger in magnitude.

Specification (5) interacts the week fixed effects with year fixed effects. This allows us to control for specific market characteristics in any given time period. For example, if a very unique movie were released in a given week across many theaters, this would account for the fact that customers with demand for this movie may be systematically different than customers arriving in other weeks. Once again, the estimated effects only become stronger. The final set of fixed effects is added in specification (6). We interact the theater fixed effects with quarter and year fixed effects. This controls for factors specific to a given theater within a time period. One advantage of this is that it can account for theaters periodically increasing prices to keep up with inflation. The results are also robust to this specification.

Table 3 describes specifications accounting for the potentially confounding factors, such as concession lines being longer when attendance is greater. In (1), we drop all observations in which the attendance for the week is greater than the average attendance at the theater. This removes occasions when lines should be longest (i.e. the highest demand weeks). In this sample of 3,524 theater weeks, we see that the relationship still holds. In (2), we include a variable that measures how much actual demand differed from what the theater forecasted it to be the week before. Such forecasts are used for staffing purposes, such that concession line length should be correlated with how far actual demand differs from forecasted demand. This variable is not significant and does not alter the relationship between concession sales and attendance. We have also tried including the forecast error in percentage terms and including the forecasted attendance in logs and neither alters the coefficients of interest.

Column (3) further explores the robustness to queuing and other confounding factors by interacting the coefficient of interest, log(Attendance) with deciles of the attendance distribution at the theater. We see that the coefficient is not significantly different than the 40-50 percent decile (which is excluded) except for the top decile in which the coefficient is 0.03 lower. This likely picks up the effect of line length resulting from fixed inputs such as soda machines rather than the variable inputs such as staffing that we proxy for with forecasted attendance. The notion is that when the theater is very busy, there may not be any level of variable inputs that can avoid long concession lines. The encouraging factor about this is that it picks up an additional drop in concession sales per person in these high attendance weeks without washing out the affect across all other levels of attendance.

Specification (3) is also useful because it narrows the scope of any factor that could confound our estimated relationship. It essentially suggests that whatever confounding factor might exist, it must be equally relevant at all attendance levels. This removes the possibility that our findings reflect systematically different types of movies with different concession demand across broadly different levels of attendance. Even within a decile of attendance, the variation in attendance reflects a negative relationship with concession demand per person. The positive relationship between willingness to pay for admission and concession demand is exactly the phenomenon which can explain this within decile relationship.

Finally, specification (4) controls for weather which also can affect affect demand for concessions. The only weather variable which has a significant effect is the average temperature during the summer. It appears that consumers might be consuming more cold beverages, for instance, on hot summer days than cooler summer days. This also does not alter the estimated relationship between concession sales and attendance.

The relation between Concession Sales and Movie Types

While our results in Table 3 suggest that estimates are not confounded by other factors, we verify this by also analyzing the characteristics of movies at the theater, which we observe during the first 26 weeks of the general sample. Table 4 shows results of five different regressions using the weighted average movie characteristics (genre, rating classification, US experience and weeks since release) in a given week at a given theater. Column (1) replicates the regression in Table 2 column (6) using theater and week fixed effects. From the results in column (1) we observe that log of attendance is still significantly less than 1 and therefore the marginal consumer left outside the theater values concessions less than the average consumer inside the theater. This result holds in columns (2) to (5) when we control by movie composition in each theater in any given week.

Column (2) replicates the regression in column (1) adding genres present in each theater. Science Fiction, Comedy, and Animated seem to have larger concession spending than the excluded genre, Fantastic. Column (3) replicates the exercise of column (2) but controlling for rating classification. We see that All Audience and PG 13 movies have lower concession spending than the excluded group, PG 7 movies. In column (4) we combine these variables into one regression and find effects for a subset of the characteristics with effects in (2) and (3). The relationship of interest remains significant throughout.

Specification (4) in Table 4 adds some additional variables. Weeks After Release, Fraction in Opening Week and US Box Office Revenue are intended to capture the phenomenon that there could be different types of customers showing up in opening weeks than non-opening weeks. None of these variables is significant and the estimated relationship between concession sales and attendance does not change. This final specification also includes weighted average US box office revenue of the movies, which also does not affect concession revenue.

The specifications throughout Tables 2, 3, and 4 account for most factors that could confound the relationship between concession revenues and attendance. The outstanding result is that when marginal customers are lured into a theater (i.e. attendance increases), the average revenues from concessions decreases. This indicates that these marginal customers consume fewer concessions, which is the necessary condition identified by Rosen and Rosenfield (1997) and Schmalensee (1981) to justify charging a premium on concessions to price discriminate.

5. Conclusions

In this paper we use a new and unique data set of weekly concession sales, box office revenues and theater attendance from a large Spanish exhibitor to examine whether theories of metering price discrimination can explain high concession prices. Even though the theoretical literature has provided many examples of industries where such types of price discrimination may occur (sport tickets and concessions within a stadium, razors and blades, printers and cartridges, Polaroid cameras and film, etc), the empirical literature has failed to provide empirical demand analyses to test the theory. We test this using one of the most widely cited examples, the sale of movie tickets and concessions in movie theaters.

Our simple empirical analysis is enabled by identification arising from a demand shock that is specific to the primary good, common to all consumers and uncorrelated with the aftermarket good. The demand shock lures (or detracts) marginal consumers from attending, thereby allowing us to infer variation in willingness to pay across customers from variation in attendance. Correlating attendance with concession sales per attendee then reveals the joint distribution of concession sales and attendance to evaluate whether demand conditions support the profitable use of metering price discrimination. The identification approach is not specific to the movie exhibition industry and can therefore be applied in other primary and aftermarket goods applications in which primary good demand varies over time due to demand shocks that are exogenous to individuals' tastes for aftermarket goods.

Our results show that as attendance increases, concession sales per attendee decreases. Our theoretical framework and set of controls in our regression estimates link high attendance weeks with customers who have lower innate tastes for movies, such that a positive correlation between attendance and concession sales per attendee results. This supports theaters' use of high margins on popcorn and other concession products to engage in metering price discrimination. This result is robust to the inclusion of controls that may reflect changes in demand across weeks, theater locations and movie types.

In summary, understanding the relative aftermarket good demands for customers with high and low values for primary goods can justify the use of metering price discrimination. In cases where this relationship does not exist, other explanations for apparent aftermarket price premiums may exist or firms may be better off lowering their aftermarket prices.

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Table 1 A
Summary Statistics

Variable	Obs.	Mean	Std. Dev.	Min	Max
Box Office Data					
Concession Sales pp	6,206	1.59	0.29	0.24	2.94
Box Office pp	6,206	4.68	0.59	2.60	6.26
Weekly Attendance	6,206	8864.27	5698.95	348	40303
Theater Characteristics					
No. Screens	43	9.65	5.20	1	24
No. Seats	43	2344.86	1248.13	396	5300
Other Variables					
Forecast Error	4,024	-652.37	2109.63	-18182	11405
Average Temperature	6,117	60.15	11.49	33.86	92.29
Rain Days	6,120	1.62	1.64	0	8

Note: Forecast Error is equal to the Weekly Attendance minus the theater's week ahead forecast of the attendance. There are only 4,024 observations for this variable because we do not observe the forecasts for the first 62 weeks of the data.

Table 1 B
Summary Statistics for Sample when Movie Characteristics are Available

Variable	Mean	Std. Dev.	Min	Max
Concession Sales pp	1.41	0.26	0.44	2.90
Box Office pp	4.32	0.48	2.70	4.98
Weekly Attendance	8522.99	5932.78	408	37565
No. Screens	7.64	3.21	2	16
No. Seats	1849.28	771.21	396	3875
Genre, Classification and Weeks After Release (Weighted by Spanish Box Office Revenue of Each Movie)				
Action	0.06	0.08	0.00	0.41
Adventures	0.34	0.24	0.00	1.00
Science-Fiction	0.14	0.23	0.00	0.96
Comedy	0.08	0.12	0.00	1.00
Animated	0.14	0.14	0.00	0.93
Drama	0.13	0.14	0.00	1.00
Fantastic	0.02	0.04	0.00	0.57
Terror	0.01	0.02	0.00	0.39
Thriller	0.10	0.14	0.00	1.00
PG13	0.40	0.21	0.00	1.00
PG18	0.10	0.13	0.00	1.00
PG7	0.06	0.09	0.00	1.00
All Audiences	0.43	0.21	0.00	0.96
US Box Office Revenue	185.75	64.07	0.00	390.06
Weeks After Release	6.21	2.84	1.00	14.89
Fraction Opening	0.13	0.17	0.00	1.00
Number of Openings*	2.04	1.32	0.00	7.00

Note: This table describes summary statistics for a sample of weekly theater obs for which movie screening are available. This sample is made of 622observations that cover the first 26 weeks of 2002. The sample contains information on 24 different theaters, we observe the complete time series for all except one.
* Number of openings is the only movie characteristic that is not a weighted average of the Spanish Box Office revenue.

Table 2
Relationship Between Concession Revenues and Attendance

Dependent Variable: log (Concession Revenue)

	(1)	(2)	(3)	(4)	(5)	(6)
log (Attendance)	1.080	0.996	0.961	0.913	0.866	0.848
	(0.004)***	(0.005)	(0.004)***	(0.004)***	(0.005)***	(0.007)***
Box Office Revenue	0.023	-0.005	0.106	0.114	0.122	0.142
per Attendee	(0.004)***	(0.004)	(0.005)***	(0.004)***	(0.008)***	(0.013)***
No. Screens		0.029				
		(0.001)***				
No. Seats per Screen		0.00008				
		(0.00004)*				
Constant	-0.374	0.235				
	(0.033)***	(0.040)***				
Fixed Effects						
Week	No	No	No	Yes	No	No
Week * Year	No	No	No	No	Yes	Yes
Quarter * Year * Theater	No	No	No	No	No	Yes
Theater	No	No	Yes	Yes	Yes	Yes
Observations	6,206	6,206	6,206	6,206	6,206	6,206
R-squared	0.94	0.95	0.98	0.99	0.99	0.99

Standard errors in parentheses. * significant at 10%; ** significant at 5%; * significant at 1%**
***'s indicate significance from zero for all variables, except log(Attendance) which is difference from 1**

Table 3
Relationship Between Concession Revenues and Attendance
Robustness to Queing and Other Confounding Factors

Dependent Variable: log (Concession Revenue)

	(1)	(2)	(3)	(4)
log (Attendance)	0.881	0.853	0.848	0.847
	(0.018)***	(0.018)***	(0.029)***	(0.029)***
Box Office Revenue	0.119	0.151	0.145	0.145
per Attendee	(0.021)***	(0.027)***	(0.025)***	(0.025)***
Attendance - Forecasted Attendance		1.10	0.76	0.66
(in millions)		(1.04)	(1.02)	(1.04)
log(Att) * Percentiles of Att				
less than 10 percentile			-0.011	-0.009
			(0.011)	(0.011)
10 to 20			0.008	0.010
			(0.009)	(0.009)
20 to 30			-0.008	-0.006
			(0.010)	(0.010)
30 to 40			-0.001	0.001
			(0.010)	(0.010)
50 to 60			0.012	0.016
			(0.011)	(0.012)
60 to 70			0.017	0.020
			(0.016)	(0.016)
70 to 80			0.003	0.003
			(0.012)	(0.012)
80 to 90			-0.008	-0.007
			(0.018)	(0.018)
90 to 100			-0.031	-0.029
			(0.011)***	(0.012)**
Number of Days with Rain				0.000
				(0.001)
Average Temperature				0.000
				(0.001)
Number of Days with Rain * Summer				0.001
				(0.002)
Average Temperature * Summer				0.002
				(0.001)***
Fixed Effects				
Week * Year	Yes	Yes	Yes	Yes
Quarter * Year * Theater	Yes	Yes	Yes	Yes
Theater	Yes	Yes	Yes	Yes
Percentiles of Attendance	No	No	Yes	Yes
Observations	3,524	4,024	4,024	3,946
R-squared	0.90	0.95	0.95	0.95

Standard errors in parentheses. * significant at 10%; ** significant at 5%; * significant at 1%**
***'s indicate significance from zero for all variables, except log(Attendance) which is difference from 1**

Table 4
Relationship Between Concession Revenue and Attendance
Accounting for Movie Genre, Classification and Weeks After Release

Dependent Variable: log (Concession Revenue)

	(1)	(2)	(3)	(4)	(5)
log (Attendance)	0.858	0.874	0.858	0.878	0.870
	(0.052)***	(0.053)**	(0.052)***	(0.054)**	(0.053)**
Box Office Revenue per Attendee	0.321	0.279	0.302	0.272	0.271
	(0.139)**	(0.129)**	(0.131)**	(0.126)**	(0.125)**
Characteristics Weighted by Spanish Box Office Revenue of Each Movie					
Action		0.240		0.287	0.263
		(0.235)		(0.252)	(0.242)
Adventures		0.223		0.339	0.305
		(0.136)		(0.137)**	(0.169)*
Science Fiction		0.269		0.232	0.230
		(0.121)**		(0.141)	(0.153)
Comedy		0.432		0.432	0.424
		(0.155)***		(0.149)***	(0.164)***
Animated		0.272		0.286	0.281
		(0.148)*		(0.181)	(0.194)
Drama		0.046		0.144	0.126
		(0.170)		(0.164)	(0.171)
Terror		0.202		0.262	0.213
		(0.191)		(0.193)	(0.189)
Thriller		0.138		0.188	0.160
		(0.136)		(0.134)	(0.141)
PG 13			-0.285	-0.187	-0.179
			(0.087)***	(0.088)**	(0.090)**
PG 18			-0.050	0.086	0.087
			(0.102)	(0.104)	(0.099)
All Audiences			-0.171	-0.051	-0.065
			(0.076)**	(0.108)	(0.126)
Weeks After Release					21.600
(in millions)					(18.400)
Fraction in Opening Week					0.004
					(0.004)
US Box Office Revenue					-0.00002
					(0.0003)
Number of Openings					-0.013
					(0.009)
R-squared	0.99	0.99	0.99	0.99	0.99

Note: All regressions contain 622 observations and use theater and week fixed effects. We drop 2 variables to avoid multicollinearity: Fantastic and PG 7 movies.
Standard errors are in parentheses. * significant at 10%; ** significant at 5%; * significant at 1%**