Hedonic Price Functions and Market Structure: An Analysis of Supply-Motivated Submarkets for Salmon in California

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ABSTRACT

Hedonic price functions are frequently used to identify the marginal implicit prices for characteristics of differentiated products, including seafood products. These marginal implicit prices are determined by both consumer preferences and producer costs. When preferences or costs for a single characteristic differ systematically across omitted submarkets, the results are an average of the heterogeneous effects. An empirical test for the joint significance of interaction terms is sufficient to identify the presence of submarkets. We motivate the selection of submarkets on the basis of production differences across salmon species, then use county-level aggregate retailer scanner data for the US state of California from 2013 to 2016 to estimate hedonic price functions. The results indicate that species and production method submarkets exist. Letting differences in production motivate the dimensions for submarkets results in a more accurate picture of the market.

Key words: Aquaculture, fisheries, hedonic prices, market structure, retail scanner data, salmon.
JEL codes: C23, C52, D49, L11, Q11, Q22.

INTRODUCTION

Using the broadest definition, the market for seafood consists of a wide variety of species with a number of diverse characteristics. However, with the variation in production costs and consumer preferences across these products, identifying submarkets becomes an important step in economic analysis. Market integration and demand studies frequently use the generalized composite commodity theorem to identify market segments or the lack thereof (Asche, Bremnes, and Wessells 1999; Asche and Wessells 2002; Xie and Myrland 2011; Salazar and Dresdner 2021). For salmon as well as other major species groups, the results largely agree that there is a single global market for each species group with a common underlying price determination process, while hedonic price studies identify significant price heterogeneity within species groups that allows producers to benefit from product differentiation (Pettersen and Asche 2020). Unfortunately the role of production costs in this price heterogeneity has been largely ignored when estimating hedonic price functions

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for seafood, with empirical papers often estimating one hedonic price function for multispecies data (Roheim, Gardiner, and Asche 2007; Sogn-Grundvåg, Larsen, and Young 2014; Bronnmann and Asche 2016).

Because the hedonic price function is the joint envelope of the consumers' value (or bid) functions and the producers' offer functions (Rosen 1974), estimates of this function may be misleading if either consumer preferences or production costs for an attribute differ systematically over species, production method, or other dimensions. In the context of residential housing, Straszheim (1974, 405) writes, "If the true structural coefficients for the several [attributes] vary across submarkets, there is no substitute for stratifying the data before estimation, or otherwise allowing these particular coefficients to vary across submarkets." A specific example for ecolabeled seafood would be if the cost of using sustainable practices varied by species, or if the consumers' interest in ecolabels varied by species, then estimating each species separately or using interaction terms is the only way to get a complete picture of the market. Interaction terms are used to account for such differences in Asche and Bronnmann (2017), where they report a statistically significant 7.5% premium for the Marine Stewardship Council (MSC) label in the pooled model. However, by species the premium is –1% for saithe, 3.7% for Alaska pollock, and 30.5% for cod. While the 7.5% premium represents an average price premium across the species, it masks the economically significant variation across species. The combination of stratified models and interaction terms is used in estimating a hedonic price function for Norwegian ex-vessel cod prices in Pettersen and Asche (2020), where separate regressions are estimated for frozen and fresh cod and regional effects are estimated with interaction terms. For every characteristic across which they allow heterogeneous effects, significant differences are observed. The choice experiment study by Nguyen et al. (2015) concludes that consumers' values for characteristics such as product form, production method, and country of origin do significantly differ across species. Additionally, Hukom et al. (2020) estimates separate hedonic price functions for wild-caught cold-water shrimp and farmed warm-water shrimp, revealing significant differences in the characteristics' implicit prices across this production/species dimension. Investigating beef and pork, Parcell and Schroeder (2007) estimate separate models not just for species, but for different cuts like steaks, roast, chops, and the like, and find that the coefficients vary, suggesting the presence of submarkets. Given the variation in production costs over various species and production methods on the supply side, and the likely differences in consumer preferences, it seems clear that submarkets should receive greater attention in estimating hedonic price functions for seafood.

A procedure to test for the presence of statistically significant submarkets becomes apparent when the issue is seen as an omitted variable bias, with the submarkets associated with interaction terms omitted from the model. Two conditions are required for omitted variable bias to occur: correlation between the included variables and the omitted variables, and a nonzero effect of the omitted variables. Interaction terms are correlated with the included variables by construction, so we propose the use of tests of the joint significance of the interaction terms to identify the presence of submarkets. If the test indicates statistically significant submarkets, one should evaluate the economic significance based on the magnitude of the estimated interaction term coefficients and knowledge of the market structure.

Using retail sales of fresh and frozen salmon in California, we estimate hedonic price functions and explore the presence of submarkets by testing the joint significance of interaction terms. The results indicate that the marginal implicit prices estimated by the pooled model mask significant heterogeneity in the market. Such differences are due to some combination of consumer preference differences across species and variation in production costs resulting from such factors as species abundance and catch limits, the timing and length of their annual runs, or the difficulty of farming. Such supply-side factors can be expected to be present in data for other seafood products, and even in other agricultural products including meat and produce. Our procedure to test for submarkets in hedonic price functions, as well as the logic of using production cost variation to identify possible submarkets, extends to many differentiated agricultural products.

Additionally, we highlight the usage of analytic weighting for the estimation of hedonic price functions using retailer scanner data. Each observation is an aggregation of sales, ranging from less than a pound to over 20,000 pounds,[1] and the impact of each observation in estimation should be appropriately weighted. Analytic weighting is preferred to frequency weighting here because it accounts for the aggregation in computing standard errors, and the allowance for non-integer weights better fits our data that include variable-weight products sold at the meat counter (Dupraz 2013). These weights produce identical point estimates to the more familiar frequency weights, but produce more conservative standard errors in accordance with the uncertain distribution about the group average. As an example, for their study using retailer scanner data Mullally and Lusk (2017) use the quantity of eggs sold as analytic weights in their estimation of the change in egg prices due to a regulatory change. This approach allows them to recover the change in average price over dozens purchased rather than the change in average price over the products uniquely identified by universal product codes (UPCs) that ignore package size and market share.

This article further contributes to the empirical literature of hedonic price function estimation for salmon by providing the first estimates from the US market. It is a significant market, with an estimated per capita consumption of 2.4 pounds of salmon per year, approximately two-thirds of which is imported (Shamshak et al. 2019). Most imports are farmed Atlantic salmon from Europe and Chile, and therefore the results will complement what is known about the Japanese and European markets. Preferences for domestic seafood, as documented for the US in Garlock et al. (2020), make results from the US market a particularly valuable addition.

The remainder of the article is laid out as follows: first the related literature is discussed, and we describe the dataset. Then the hedonic price model and model selection issues are explained. The results of the hedonic price models are next, followed by concluding remarks.

RELATED LITERATURE

Hedonic price analyses are used to estimate the marginal value of a wide range of product characteristics that can provide insights into consumer preferences and the value of a number of supply-side decisions. For example, hedonic price studies can inform the efficacy of marketing and information campaigns or the potential returns to more costly but more environmentally friendly production. In discussing the market for seafood characteristics, Wessells (2002) predicts that salmon would soon be differentiated by numerous characteristics such as production method (farmed or wild-caught), species, point of origin, and sustainability, and that consumers may have a willingness to pay for these characteristics. But in addition to consumers' willingness to pay, the hedonic price analysis also measures the production costs associated with these characteristics. This point is summarized by Costanigro and McCluskey (2012, 157) who write that "high implicit prices

1. The results are stable across models estimated with lower bounds ranging from 1 ounce to 16 ounces. Allowing for these weights between 0 and 1 does not significantly impact the results.

may very well be due to elevated costs of production, and it is possible that only a small fraction of consumers actually purchase bundles containing that expensive attribute." Although the hedonic price method has been used many times to study seafood markets, the importance of production costs has been largely ignored, especially as it relates to submarkets.

The salmon species group consists of multiple species, including the six analyzed in this study: Atlantic (*Salmo salar*) and five species of Eastern Pacific salmon, namely, Chinook (*Oncorhynchus tshawytscha*), Coho (*O. kisutch*), Sockeye (*O. nerka*), Pink (*O. gorbuscha*), and Chum (*O. keta*). Atlantic salmon are presently only commercially available through aquaculture, while Pacific salmon sold in the US are primarily from wild-capture fisheries, with only Chinook and Coho being farmed commercially during the study period. Several studies test for market integration of the species and use the generalized composite commodity theorem of Lewbel (1996) to test whether the law of one price holds, which indicates perfect substitution among the species (Asche, Gordon, and Hannesson 2004). These studies find that there is one market for all salmon species and that they are perfect substitutes (Asche, Bremnes, and Wessells 1999; Asche and Wessells 2002; Asche et al. 2005). However, while this finding is sufficient for aggregation in demand analysis, it may not extend to the estimation of hedonic price functions (Pettersen and Asche 2020). In this context, the law of one price implies a relationship between the dependent variable (price) across species; the independent variables (product characteristics) and their coefficients (marginal implicit prices) are free to differ. In order for hedonic price functions to differ across submarkets, it must be that "the structure of demand, the structure of supply, or both must be different across segments" (Freeman, Herriges, and Kling 2014, 325).[2] While it is certainly possible that the structure of demand differs across salmon species, there is no doubt that the structure of supply differs. Atlantic salmon production has fish feed as the primary input, and it is regulated through limits on the number and size of aquaculture operations in each country. Wild-caught Eastern Pacific salmon have fuel and labor as primary inputs, and they are subject to species-specific harvest regulations set by the US and Canadian governments. In spite of these differences, several previous hedonic price studies for seafood have treated the species group as a single homogeneous product (Roheim, Gardiner, and Asche 2007; Bronnmann and Asche 2016). Sometimes Atlantic salmon and the five Eastern Pacific salmon are lumped into several categories, such as in Asche et al. (2015). This generalization may not be cause for much concern in the European markets where Atlantic salmon dominates.[3] However, it is important to consider the possibility of differences in the marginal implicit prices across the species in the American and Japanese markets, where the mix of species is diverse enough that meaningful species differences would be masked by a single pooled hedonic price function.

The issue of market segmentation/submarkets can also be seen in the literature that estimates hedonic price functions for wine. There the issue has received more consideration, with Thrane (2004, 124) arguing in favor of separate hedonic price regressions for white and red wine, as "no theoretical or common-sense justification has yet been provided for why the effects of a set of attributes on wine prices are similar for red and white wines." A further segmentation of the

2. It is also worth considering geographic submarkets. Freeman, Herriges, and Kling (2014) provide a second condition, which is that arbitrage is limited by barriers to consumers or products moving to different markets. Many California counties are large enough to limit consumer movement across counties, and statistical tests identify differences in the coefficients across counties. This is beyond the scope of this paper but may be a fruitful area for future research.

3. Asche et al. (2015) report that 87.4% of the observations are Atlantic salmon.

wine market is proposed in a study by Costanigro, McClusky, and Mittelhammer (2007), which finds differences in the marginal implicit prices across different price classes of red wine. However, even in this literature the arguments for segmentation have centered on differences in preferences, with little attention given to possible differences in offer curves across submarkets.

Returning to the market for seafood, a common thread through both the stated and revealed preference literature is a preference for local/domestic seafood, a summary of which is presented in Rickertsen et al. (2017). Preference for local seafood may be attributable to perceptions of quality, freshness, or a reduced carbon footprint. Such preferences have been observed among US seafood wholesalers (Garlock et al. 2020), and would suggest that Californians prefer wild Alaskan salmon varieties and Europeans prefer farmed Atlantic salmon.[4] However, trade statistics show significant exports of domestic salmon and imports of Atlantic salmon (NMFS 2019), suggesting that the US salmon market may be an exception to this rule. There have been no hedonic price studies of salmon in the US, but there are studies estimating hedonic price models for salmon purchased in the United Kingdom (Asche et al. 2015), Denmark (Ankamah-Yeboah, Nielsen, and Nielsen 2016), and China (Du et al. 2020). The market integration analysis of Salazar and Dresdner (2021) finds that the law of one price holds for farmed Atlantic salmon of different origins sold in the US. This still leaves room for constant price differences that can be attributed to transportation costs or quality differences that would manifest in a hedonic price function. Previous hedonic studies of salmon prices have given little attention to the point of origin, and the issue is complicated by correlation between origin and species. The only estimates of the impact of origin for salmon are in Asche et al. (2015), where UK consumers are found to have a positive but insignificant premium for Scottish-origin salmon, and a negative but insignificant premium for Alaskan-origin salmon.

A major focus for the existing retail seafood demand literature has been the price premium for ecolabels, such as MSC certification (Roheim, Asche, and Santos 2011; Sogn-Grundvåg, Larsen, and Young 2013, 2014; Blomquist, Bartolino, and Waldo 2015; Asche and Bronnmann 2017), organic labels on farmed salmon (Ankamah-Yeboah, Nielsen, and Nielsen 2016; Ankamah-Yeboah et al. 2020), and Aquaculture Stewardship Council certification (Bronnmann and Asche 2017). The aforementioned studies find that these credence attributes are associated with a roughly 10%–20% price premium.

Some of the ecolabel studies also distinguish between farmed and wild-caught as an ancillary variable in their analyses, with conflicting results across the literature. Two closely related stated preference articles, Holland and Roheim (1998) and Roheim, Sudhakaran, and Durham (2012), estimate preference rankings for farmed and fresh salmon among consumers in New England. The earlier article reports that farmed is preferred to wild-caught (with the caveat that they expect the results would be different on the West Coast), but the later article finds that wild-caught is preferred. This change could be the result of changes in preferences or product quality, or it could be related to increasing negative publicity regarding the environmental impacts of salmon farming and health risks of consuming farmed salmon (Amberg and Hall 2008). The French stated preference studies of Nguyen et al. (2015) and Rickertsen et al. (2017) also find a preference for wild-caught over farmed for a number of different species. In contrast to these, the revealed preference study of Bronnmann and Asche (2016) finds a positive price premium on farmed seafood in the

4. Alaska is both the closest source of salmon for the California market and in the same country. However, it is also a significant distance from California so it is not clear to what degree Alaskan salmon would achieve the "local premium" that is generally observed in food demand.

German frozen seafood market. In order to gain insight into the source of the wild-caught premium, Bronnmann and Asche (2017) use a stated preference survey to evaluate preferences for wild-caught and farmed salmon with and without ecolabels. The results suggest that consumers are concerned about sustainability, as the positive effect of the Aquaculture Stewardship Council label overcomes the negative premium for farmed salmon. Market integration studies show more consistency, with Asche, Bremnes, and Wessells (1999) and Asche et al. (2005) finding that there is one market for salmon, irrespective of the species or production method. Xie and Myrland (2011) report that wild-caught and farmed salmon compete in different niches, but they have made this assertion on the basis of differences between canned and non-canned salmon. This lack of market integration may be a result of, for instance, shelf stability, perceived quality differences, or imperfect substitutability, rather than the production method.

The value of supermarket scanner data for estimating demand and market power was extolled by Cotterill (1994), but these data are less frequently used for seafood hedonic price analyses. Supermarket scanner data have been used in Roheim, Gardiner, and Asche (2007) and Roheim, Asche, and Santos (2011) to estimate hedonic price models for seafood in the United Kingdom. Supermarket scanner data typically contain the weekly quantities and sales weighted average price (total revenue divided by total quantity) aggregated by retailer or by geographic market. This average price folds into it both spatial and temporal variation in price, as it is likely to be aggregated across midweek price changes and across differently priced stores on the same day. The quantity-weighted average will more accurately represent consumers' willingness to pay, as it encapsulates their choice of retailer and promotional timing. This is part of the trade-off compared with in-person observation of product characteristics and prices, as the data on product characteristics are less accurate/complete, but by including more retailers and time periods as well as sales data, the quantity-weighted average price accounts for realized consumer behavior and not just retailers' expectations. This averaging of the price across each unit of observation demands that both the coefficient estimates and the standard errors should account for this data structure, which can be accomplished by the use of analytic weights such as in Mullally and Lusk (2017).

DATA

For our analysis, we use retailer scanner data provided by FreshFacts, a division of Nielsen focused on fresh foods such as the meat counter and produce section. The data contain the weekly purchases of raw salmon (both packaged and unpackaged), aggregated at the county level for the state of California from January 2013 to December 2016.[5] For each county-week there is an entry for each product identifier, analogous to a UPC.[6] For each product identifier, the data include information on brand, origin, form, species, production method, and a 30-character (truncated) product description.

The price is computed as the county-week revenue (US dollars) divided by volume (pounds), for each product. This means that in most cases the price has been averaged over multiple stores with different prices and over multiple pricing regimes within a store.[7] This quantity-weighted

5. Of the 58 counties in California, six of the rural counties had no retailers contracted with Nielsen and are therefore not included in the data.

6. A variable for UPC code is present in the data but does not match any known UPCs in web-based UPC databases.

7. Three of California's major grocers (Ralphs, Vons, and Safeway) change prices weekly on Wednesdays, while the data are from Sunday to Saturday. Walmart changes their prices on the first Wednesday of each month.

average price has the advantage of accounting for consumers making more purchases at the stores and times with better prices, rather than treating each observed price with equal weight. Retailers may also differ in both their base price and their hedonic coefficients as shown in Asche et al. (2015), but in the absence of retailer-specific information the sales-weighted price included here should be effective at estimating an average consumer willingness to pay.

Some characteristics previously used in the literature are not available in the data, such as package size and ecolabeling information. With a mixture of variable-weight and fixed-weight products, package size is less useful as an explanatory variable. Whether the product is fresh, frozen, or previously frozen is not in the raw data but can be inferred from the product description for 34% of observations. It is unfortunate that information about ecolabels is not available, as these are known to command a premium, but the status of ecolabels and salmon during the sample period of 2013–16 is quite complex. The Alaskan salmon fisheries dropped out of the MSC recertification process in 2012 to set up their own ecolabel (Foley and Hébert 2013), only to return to MSC certification in 2015. The Alaskan-origin dummy should capture these effects regardless, as the entire fishery (all species and gear types) went through this back-and-forth certification process at the same time. The Aquaculture Stewardship Council certified the first salmon farms around the start of our sample in 2014 but may have needed time to build credibility with consumers and to develop enough supply to export to the California market. Some of this impact is likely to be captured by the brand fixed effects, lessening any omitted variable bias.

The raw data include a number of inconsistencies, contradictions, and missing characteristics that needed to be addressed. The full description of the data cleaning and imputation process is included in the online appendix. Trimming the data is relatively common when working with scanner panels, such as Roheim, Asche, and Santos (2011), who removed all products that appear in less than one-third of weeks during the sample. However, pruning the data can introduce bias (Andrews and Currim 2005), so we endeavored to keep as much data as possible and minimize the impact of low-volume outliers through analytic weighting. Because fresh Pacific salmon is a seasonal product, no restrictions were placed on the number of weeks a product needed to be in the data. We removed 12 observations due to unrealistic prices under \$0.15 per pound or over \$100 per pound. We also removed 131,023 observations for which species was unavailable, as this is a critical characteristic for the analysis.

The cleaned dataset includes 396,926 observations totaling 45,912,933 pounds and \$377,174,723 over the four-year sample. Not surprisingly, the price for wild-caught salmon is more volatile than that for farm-raised salmon (see online appendix figure A1), in agreement with Dahl and Oglend (2014) and Asche, Dahl, and Steen (2015). Because the data lack evidence of an inflationary trend, the nominal price is used in conjunction with time fixed effects in all models.

The descriptive statistics in table 1 show a relatively large range of prices across most categories. Because of these wide ranges, the discussion that follows centers on the means, frequencies, and market shares. The ranking of species generally aligns with previous findings such as Asche et al. (2005) and Asche et al. (2015), with Pink and Chum as low-value salmon species, Atlantic and Coho in the middle,[8] and Sockeye and Chinook (often marketed as King salmon) as high-value ones. Atlantic and Sockeye make up the majority of the market share, with Chum, Coho, Chinook,

8. Asche et al. (2005) find that Atlantic salmon is a high-value species, but Asche et al. (2015) report that the price of Atlantic salmon is less than Sockeye and is comparable to “Wild Alaskan” salmon.

Table 1. Descriptive Statistics of the Product Characteristics

Product Attribute	Frequency (%)	Mean Price ($/lb.)	Min. ($/lb.)	Max. ($/lb.)	Std. Dev. ($/lb.)	Volume Share (%)	Market Share (%)
Species							
Atlantic	42.6	8.41	0.19	34.78	2.75	71.7	67.8
Sockeye	32.0	12.77	0.67	49.98	3.75	14.5	19.5
Chum	9.9	7.33	0.49	21.28	2.24	6.2	4.1
Coho	7.0	9.36	0.50	30.35	3.74	3.7	3.8
Chinook	6.1	15.81	0.59	53.99	6.00	2.6	4.3
Pink	2.3	5.90	0.86	11.35	2.03	1.3	0.6
Product form							
Fillet	62.2	10.73	0.49	53.99	4.21	78.7	82.3
Portion	14.1	9.68	0.79	48.07	3.73	9.5	7.8
Unknown	10.2	10.38	0.50	49.98	5.09	6.6	6.7
Steak	9.6	8.86	1.79	38.99	2.97	2.1	2.1
Whole	3.2	5.91	0.79	30.99	3.17	2.9	1.1
Heads	0.6	2.24	0.19	7.99	1.09	0.1	0.0
Origin							
Generic	59.1	8.19	0.19	34.78	2.93	80.2	72.8
Alaska	36.4	13.02	0.59	49.98	4.20	15.4	21.4
Copper Riv.	1.9	16.13	0.67	53.99	5.73	1.1	1.8
Norway	1.6	10.31	3.33	21.50	1.36	3.1	3.7
Chile	1.0	10.27	2.00	15.12	3.18	0.2	0.3
Scotland	0.0	16.36	11.30	20.06	2.42	0.0	0.0
Production method							
Wild	55.3	11.53	0.49	53.99	4.70	27.3	31.2
Farmed	44.7	8.46	0.19	34.78	2.84	72.7	68.8
Condition							
Unknown	59.3	9.95	0.50	53.99	4.12	60.2	60.6
Fresh	22.4	10.40	0.19	40.00	5.01	30.4	29.8
Frozen	11.6	10.38	0.75	26.99	3.71	6.9	7.0
Prev. frozen	6.7	10.82	0.49	25.99	3.49	2.4	2.6

Note: $N = 396{,}926$.

and Pink salmon combining to total about 10% of the market. The product-form statistics also conform with expectations about consumer preferences, with the time-saving processed cuts of fillet, steak, and portion (typically fillets cut down to individual serving sizes) having higher prices while the more labor-intensive whole fish are cheaper. Heads and bones (simply "heads" in tables), which are most commonly used for soup, are less expensive. Fillets are the most popular product form by far, holding 82.3% of the market by revenue. Country of origin labeling has been mandatory for salmon sold in the United States since 2005, but surprisingly the generic-origin salmon makes up 72.8% of the market in the data. This would suggest that consumers had this information while we do not, although consumer inattention to a small-print origin identification is also probable. Therefore, our results regarding origin must be viewed with this caveat in mind. Approximately 90% of US-caught Pacific salmon comes from Alaska (NMFS 2019), so this is functionally equivalent to a US-origin label, and these products make up 21.4% of the market. Copper River, an Alaskan river known for high-quality salmon (see Jardine, Lin, and Sanchirico 2014), and Scottish origin receive a larger premium compared with Alaskan, Norwegian, and Chilean. The production method also appears to be an important feature, as on average the wild-caught salmon cost an additional $3 per pound (38% more). Intriguingly, about 45% of the products are wild-caught but only represent 31% of the market share. With respect to the condition at the time of sale, this variable could not be inferred from the product description for 59.3% observations, and there is little difference in the mean price for fresh, frozen, and previously frozen.

MODEL SPECIFICATION AND TESTING PROCEDURES

The hedonic price model assumes that, within a class of products differentiated by a number of characteristics, the price of a particular product is determined by its bundle of characteristics (Rosen 1974). In this case, those characteristics include features such as species, place of origin, product form, farmed or wild-caught, and whether it is or was previously frozen. Each of these characteristics determines the utility from consumption, or in the case of product form, the value of shifting preparation to the producer. At the same time, each of these characteristics has an accompanying cost of production. The general form of the hedonic price function in Rosen (1974), written for our panel data context, is simply

$$P_{ct}^{j}\left(X^{j}\right) = f\left(x_{1}^{j}, \ldots, x_{k}^{j}\right),$$

where P_{ct}^{j} is the price of product j in week t in county c and $X^{j} = (x_{1}^{j}, \ldots, x_{k}^{j})$ is a vector describing the k characteristics of product j that factor into the price.

No particular functional form for f is prescribed by theory or the empirical literature. The results of a Box-Cox analysis (Box and Cox 1964) for the pooled data models (table 2) and for the species-specific models (table 3) suggest a log-linear specification is appropriate. Although there would be efficiency improvements from using the exact Box-Cox transformation, a consistent log-linear specification simplifies the interpretation and comparison of the results across models.

With theory and the empirical literature indicating the likelihood of submarkets being present, we use a simple hypothesis test to evaluate possible submarket dimensions. The goal is to evaluate the difference between the baseline pooled model, estimated on the full dataset in which each product characteristic has an independent level effect, and the stratified model, estimated on data sliced across some dimension in the data so that each product characteristic is allowed to have different effects over that dimension. To facilitate the hypothesis test, we note that the stratified model is equivalent to adding a full slate of interaction terms for the stratification variable (slope effects) to the pooled model estimation, in what we refer to as the interacted model. Because the interacted model is nested in the pooled model, a number of tools are available to evaluate the differences. Both the likelihood-ratio test and the general linear F-test can easily evaluate the validity of restricting the interaction terms to zero as is assumed in the pooled model. One could also estimate the stratified models and apply a Chow test as in Pettersen and Asche (2020). However, these tests assume independent and identically distributed normal errors, and therefore are not appropriate in the presence of heteroscedastic or clustered errors, so we use the Wald test with a robust variance-covariance matrix. If the test rejects the null that the coefficients on the interaction terms are jointly zero, then there are statistically significant submarkets and the stratified or interacted models should be evaluated for economic significance.

The characteristics of interest, such as origin and production method, are included as a series of indicator variables. Additional fixed effects are included to account for other sources of price variation. County fixed effects absorb geographic variation due to grocers' overhead and labor costs, transportation, and local consumer characteristics. A set of year-month fixed effects provide the

Table 2. Hedonic Price Function Estimates with Pooled Species

	Unweighted (1)	Analytic Weights (2)
Production method	(Base: Farmed)	
Wild	0.34**	0.36**
	(0.11)	(0.10)
Product form	(Base: Fillet)	
Heads	–1.56***	–1.66***
	(0.11)	(0.03)
Portion	–0.05	–0.32***
	(0.03)	(0.05)
Steak	0.00	0.01
	(0.02)	(0.01)
Whole	–0.42***	–0.66***
	(0.10)	(0.07)
Unknown	–0.15**	–0.03
	(0.04)	(0.09)
Condition	(Base: Unknown)	
Fresh	0.10*	–0.04
	(0.05)	(0.06)
Frozen	–0.09	–0.12*
	(0.05)	(0.05)
Prev. frozen	–0.06*	–0.16
	(0.03)	(0.08)
Origin	(Base: Generic)	
Alaska	0.06	0.18*
	(0.08)	(0.09)
Chile	0.82***	0.80***
	(0.10)	(0.16)
Copper Riv.	0.38***	0.36***
	(0.07)	(0.08)
Norway	0.18***	0.14**
	(0.03)	(0.05)
Scotland	0.56***	0.44***
	(0.04)	(0.05)
Species	(Base: Atlantic)	
Chinook	0.25**	0.08
	(0.08)	(0.09)
Chum	–0.54***	–0.64***
	(0.12)	(0.12)
Coho	–0.16	–0.19*
	(0.10)	(0.09)
Pink	–0.62**	–0.92***
	(0.17)	(0.14)
Sockeye	–0.03	–0.24*
	(0.09)	(0.11)
Constant	2.06***	2.14***
	(0.03)	(0.06)
Observations	396,926	396,926
R^2	0.59	0.69
SSE	31,574	14,838
Box-Cox θ	0.31	0.16
χ^2 stat ($\theta = 0$)	15,534	2,584

Note: Cluster (species) robust standard errors are in parentheses. Model 2 employs analytic weighting on volume sold (pounds). All models include fixed effects for county, year-month, and brand name. The omitted factors are Los Angeles County, December 2016, and unbranded. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table 3. Hedonic Price Function Estimates by Species

	Atlantic (1)	Chinook (2)	Chum (3)	Coho (4)	Pink (5)	Sockeye (6)
Production method						
Wild	–0.42***			0.19***		
	(0.09)			(0.05)		
Product form						
Heads	–1.63***	–1.86***				
	(0.17)	(0.05)				
Portion	–0.37***	0.31*	0.25***	0.56***	0.11	–0.14***
	(0.03)	(0.18)	(0.06)	(0.10)	(0.17)	(0.03)
Steak	0.00	–0.07	–0.32***	0.05	0.47**	–0.10
	(0.02)	(0.08)	(0.08)	(0.04)	(0.19)	(0.08)
Whole	–0.33***	–0.42***	–0.50***	–0.75***	–0.48***	–0.64***
	(0.04)	(0.07)	(0.05)	(0.13)	(0.15)	(0.18)
Unknown	0.09**	–0.10*	–0.49***	–0.08	–0.17	–0.14***
	(0.04)	(0.06)	(0.04)	(0.09)	(0.19)	(0.02)
Condition						
Fresh	–0.12***	0.16***	0.05**	0.08	0.14	0.10***
	(0.04)	(0.044)	(0.02)	(0.06)	(0.24)	(0.03)
Frozen	–0.55***	–0.49***	–0.22***	–0.45***	0.42**	–0.10***
	(0.04)	(0.12)	(0.06)	(0.03)	(0.17)	(0.02)
Prev. frozen	–0.43***	–0.07	–0.12	–0.12	–0.09	–0.05*
	(0.07)	(0.09)	(0.11)	(0.08)	(0.12)	(0.03)
Origin						
Alaska		0.54***	0.30***	0.17***		–0.22***
		(0.08)	(0.09)	(0.06)		(0.05)
Chile	0.41***			0.69***		
	(0.03)			(0.15)		
Copper Riv.		1.21***		0.13***		
		(0.09)		(0.04)		
Norway	0.08***					
	(0.03)					
Scotland	0.56***					
	(0.07)					
Constant	2.19***	2.33***	1.83***	2.07***	1.29***	2.53***
	(0.01)	(0.08)	(0.02)	(0.05)	(0.16)	(0.04)
Observations	169,254	24,076	39,397	27,876	9,257	127,066
R^2	0.47	0.83	0.85	0.79	0.89	0.47
SSE	5,264	780	1,034	882	225	3,631
Box-Cox θ	0.19	0.09	0.19	–0.05	–0.38	–0.10
χ^2 stat ($\theta = 0$)	718	71	404	19	419	173

Note: Two-way (county by year-month) cluster robust standard errors are in parentheses. All models employ analytic weighting on volume sold (pounds). All models include fixed effects for county, year-month, and brand name. The omitted factors are Los Angeles County, December 2016, and unbranded. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

most flexible handling of variation in prices over time due to changes in preferences, variations in supply, and inflation.[9] Brand fixed effects soak up any effects due to unobserved characteristics particular to a brand. The baseline product is an unbranded fillet of farmed Atlantic salmon with

9. Less granular time fixed effects, such as month & year and quarter & year, produce similar results, but do poorly at capturing variation in 2015–16. See online appendix figure A2.

generic origin and unknown condition (fresh or frozen) sold in Los Angeles County in December 2016. The *pooled model* is specified as the following:

$$\begin{aligned}\ln\left(P_{jct}\right) &= \alpha + \sum_{c=2}^{52}\beta^{c}_{cnty}cnty_c + \sum_{ym=2}^{48}\beta^{ym}_{ym}ym_t + \sum_{b=2}^{39}\beta^{b}_{brand}brand_j \\ &\quad + \sum_{k=2}^{3}\beta^{k}_{pm}pm_j + \sum_{l=2}^{5}\beta^{l}_{pf}pf_j + \sum_{m=2}^{4}\beta^{m}_{co}co_j + \sum_{n=2}^{7}\beta^{n}_{sp}sp_j + \sum_{o=2}^{6}\beta^{o}_{or}or_j + \varepsilon_{jct},\end{aligned} \tag{1}$$

where P_{jct} is the price of product j in county c and week t, α is the average price of the baseline product, and β^{c}_{cnty}, β^{ym}_{ym}, and β^{b}_{brand} are county, year-month, and brand fixed effects, respectively, with $cnty_c$, ym_t, and $brand_j$ being identity vectors. The summations represent the deviations from the baseline for a product's production method (identity vector pm_j), product form (pf_j), condition (co_j), species (sp_j), and origin (or_j), with the summation starting at 2 to avoid the dummy variable trap and ending at the last characteristic level. Lastly, ε_{jct} is the error term.

The *interacted model* for hypothesis testing adds the interaction of the stratification variable (species) with the other characteristics (production method, product form, condition, and origin) and the fixed-effects variables (county, year-month, and brand). Switching to vector notation for convenience, the model is specified as follows:

$$\begin{aligned}\ln(\boldsymbol{P}) &= \boldsymbol{\alpha} + \boldsymbol{cnty\beta}_{cnty} + \boldsymbol{ym\beta}_{ym} + \boldsymbol{brand\beta}_{brand} \\ &\quad + \boldsymbol{pm\beta}_{pm} + \boldsymbol{pf\beta}_{pf} + \mathbf{co}\boldsymbol{\beta}_{co} + \boldsymbol{sp\beta}_{sp} + \boldsymbol{or\beta}_{or} \\ &\quad + \boldsymbol{sp} \odot [\boldsymbol{cnty\ ym\ brand\ pm\ pf\ co\ or}]\boldsymbol{\gamma} + \varepsilon,\end{aligned} \tag{2}$$

where $\boldsymbol{sp} \odot [\boldsymbol{cnty\ ym\ brand\ pm\ pf\ co\ or}]$, in a slight abuse of notation, represents the concatenated column-wise Hadamard product of each species dummy column with each other dummy column. Then the test for statistically significant submarkets is a test for $\boldsymbol{\gamma} = \mathbf{0}$.[10] While one might argue that the interaction with the fixed effects (county, year-month, and brand) is unnecessary, this is representative of the flexibility of the stratified model. One could choose not to include those interaction terms, or to test only the product characteristics coefficients; however, the omission of these variables could still be responsible for omitted variable bias on the coefficients of interest.

With data aggregated at the product-county-week level, each observation contains the total volume and total revenue from which average price is computed. As shown in table 1, the frequency with which a characteristic showed up in the data could be markedly different than the volume or revenue associated with that characteristic. Failure to account for this aggregation effect would lead to outsized impacts from infrequently purchased products and smaller counties. As an example, consider two hypothetical ecolabeled products: product A, with a premium of $1 and 100 pounds sold, and product B with a premium of $5 and 1 pound sold. Without weighting, the estimated marginal implicit price would be $3, but weighting by volume sold yields an estimated premium of $1.04. Whether one is concerned with how much the average consumer is willing to pay for the characteristic or with producers covering the increased costs, the weighted average more accurately captures this information. It is for these reasons that Mullally and Lusk (2017) weight their price

10. This model is easily described in Stata as i.sp##(i.cnty i.ym i.brand i.pm i.pf i.co i.or). The hypothesis test for $\boldsymbol{\gamma} = 0$ is executed with testparm i.sp#(i.cnty i.brand i.ym i.pf i.cond i.pm i.or).

regression by the quantity of eggs sold. The weighted regression produces different coefficient estimates as a result of minimizing the weighted sum of squares, as well as different standard errors due to changes in the model formulation. Analytic weighting is the appropriate technique when each observation is the mean of a number of transactions (Dupraz 2013). This method produces the same point estimates as the more familiar frequency weighting in which each observation represents the same values observed multiple times, but has the advantage of accommodating non-integer weights and producing standard errors adjusted for the number of observations rather than the sum of the weights, thus producing larger standard errors. In this study, pounds sold is the chosen weighting variable.

With this type of data, it is likely that standard errors will be correlated within clusters. This has been acknowledged in recent seafood hedonic studies (Asche et al. 2015; Bronnmann and Asche 2016; Asche and Bronnmann 2017). It is rarely clear how to cluster the data, and Cameron and Miller (2015) recommend simply testing ever larger clusters, choosing the smallest clustering level after which the standard errors are stable.

To this end, comparisons were made over various clustering dimensions for both the pooled and the species-specific models (see online appendix tables A1–A3) with the decision made on the basis of the observed standard errors and the underlying logic. For the pooled model, clustering on species produces larger standard errors and makes more sense given the market and producers' ability to select the species. Clustering on the species dimensions allows for the standard errors to be correlated across time (including autocorrelation), space, cut, and origin within a species, but imposes zero correlation across species regardless of time and space. This latter restriction was relaxed in the two-way clustering estimates as described in Cameron, Gelbach, and Miller (2011), but in all cases the standard errors are not appreciably different from the one-way errors. The species-specific models are clustered using two-way (county and year) clusters. This two-way cluster allows errors to be correlated across counties for a given month (e.g., supply shocks) and across time for a given county (e.g., regional preference differences or autocorrelation).

RESULTS

The pooled models results are presented in table 2, with the unweighted model in column 1 and the model weighted by pounds sold in column 2. The baseline product is unbranded generic-origin farmed Atlantic salmon fillets sold in Los Angeles County in December 2016. Both models fit the data reasonably well, with R^2 of 0.59 and 0.69, respectively. The cluster-robust standard errors are on average 15 times larger than ordinary robust standard errors (see online appendix table A1), providing evidence of clustered correlation in the errors for seafood hedonic price functions.

The use of analytic weighting results in changes to both the estimated coefficients and their standard errors, which combine to produce changes in the statistical significance of several variables. The unknown product form, fresh and previously frozen, and Chinook have a statistically significant coefficient without weighting but are not significant with the weights. Conversely, the coefficients on portion, frozen, Alaskan origin, and the species Coho and Sockeye become significant with the weighting. This underscores that failure to use weights for aggregate data can lead to meaningful changes in the implications of hedonic price function estimates. In light of previous research showing preferences for locally produced seafood, it is somewhat surprising that relative to generic-origin products, the premium for Alaskan salmon is relatively small and only significant at the 10% level.

However, these results will be an oversimplification if there are market segments that are not accounted for by stratification or interaction terms. As was discussed earlier, the joint significance of the interaction terms in the pooled model is sufficient to establish the presence of statistically significant submarkets. In the homoscedastic case, this is equivalent to the general linear F-test for reducing the sum of squared errors used in Straszheim (1974). The SSE are shown in tables 2 and 3, and comparing the pooled model of column 2 with the sum across species in table 3 produces an F-statistic of 166.2 with a critical value of 1.1 (555 numerator degrees of freedom and 396,222 denominator degrees of freedom). However, this test is not appropriate for heteroscedastic data. Instead, using the Wald test for the joint significance of the interaction terms from the model with robust standard errors,[11] the resulting F-statistic is 141.4. Both of these are well in excess of the critical value of 1.1, indicating that market segments exist. Narrowing the focus by separately testing the joint significance of species interaction terms for each set of characteristics (e.g., origin, product form, county fixed effects) confirms that the systematic difference across species in the hedonic price function is present for each characteristic set individually. Likewise, tests of each species-specific model individually against the pooled model reject the validity of aggregation at the 99.99% level. Therefore, we proceed to examine the results from the model stratified by species.

Estimating the model separately for each species allows for simpler presentation of the results compared with including hundreds of interaction terms. The results in table 3 are estimated with a baseline product of unbranded generic-origin farmed fillets sold in Los Angeles County in December 2016,[12] and demonstrate the magnitude of difference present for the marginal implicit values in each species submarket. In each of the categories of characteristics there is at least one case of an attribute having a positive and significant effect for one species and a negative and significant effect for another species, which we interpret as evidence of economic significance of the submarkets. In conjunction with the observation of different ecolabel premiums across retailers in Asche et al. (2015) and across species in Bronnmann and Asche (2016), it is clear that there exists meaningful variation in the bid and/or offer functions for seafood across a number of dimensions that warrant consideration in economic research.

The model fits the data well for the less common species, with R^2 ranging from 0.79 to 0.89 for Chum, Coho, Chinook, and Pink salmon. For the more common species of Atlantic and Sockeye, the R^2 is lower at around 0.47, suggesting the possibility that some unobserved characteristics are important for understanding the price of these species or that there is simply greater price variability.

The production method is statistically significant in all cases, with the expected positive coefficient for the pooled model and for Coho. Wild-caught Atlantic salmon was rare in the sample, with just 53 observations originating from Scotland. However, these observations are suspect, and may be mislabeled with respect to production method, species, or origin (Kroetz et al. 2020). Because there are no wild-caught Atlantic salmon from other origins, the negative coefficient is interpreted to show that what is labeled as wild-caught Scottish salmon was lower priced than farmed Scottish salmon but higher priced than farmed generic-origin salmon. Turning to Pacific salmon species and using the formula $(e^{\beta}-1) \times 100$ to compute the percentage change, the results show that wild-caught Coho earned a premium of 20.9% (0.19 coefficient). The wild-caught coefficient is not defined for Chinook because it is perfectly multicollinear with the Alaska and Copper River origins, but with these origins having premiums of 71.6% (0.54) and 235.3% (1.21) respectively,

11. The clustered covariance matrix could not be used for this test because the 675 conditions being tested exceed the number of clusters. We recommend using the clustered covariance matrix when possible, but the robust covariance matrix still produces more conservative test statistics than the homoscedastic case or the general linear F-test.

12. Not all species are farmed, thus the wild-caught coefficient is not defined for these species.

it suggests that farmed Chinook is significantly discounted in relation.[13] The estimated premium of 99.4% (0.69) on Chilean-origin Coho is a result of a collinearity issue described in the next paragraph.

Because we have argued for the stratification by species based on differences in the structure of supply, it follows logically that further stratification may be necessary for any species with both wild-capture and aquaculture. Within the models for Chinook and Coho, tests for the joint significance of the production method interaction terms reject the null hypothesis of insignificance. With critical values less than 2, the robust Wald test for Coho produces an F-statistic of 28.1, and for Chinook it is 52.7. Using the results of the stratified model (see table 4), we can still compute the price premium by a comparison of the constant (intercept) terms. Provided the excluded categories are the same, the difference between the intercepts of the two models will correspond to the coefficient on wild-caught from the pooled model. For Coho, the intercept for wild-caught is 2.52 and the intercept for farmed is 2.32, implying a coefficient of 0.20 and a percentage premium of 22.1%. Intriguingly, the effect of Chilean origin in this model is a statistically insignificant –14.8% (–0.16) compared with the 99.4% (0.69) premium from table 3. This is explained by the fact that all farmed Chilean Coho come from one low-priced brand that also offers wild-caught Coho. With both wild and farmed in the sample, the brand fixed effect is strongly negative, but the price is closer to the average for their farmed than wild-caught Coho so the estimated effect of Chilean origin is positive. When estimating on the farmed subsample, the brand fixed effect is perfectly collinear with Chilean origin, and therefore dropped from the model. The resulting coefficient on Chilean origin is pulled down as it absorbs the brand effect. Ultimately the data are unable to effectively value Chilean origin for Coho salmon. The results for Chinook not only indicate differences in the marginal implicit prices, but the results of the Box-Cox tests suggest that different functional forms may be called for, with wild-caught having an optimal θ of 0.60, and –0.07 for farmed. The data support our argument that the structure of supply is an important dimension of heterogeneity to consider in estimating hedonic price functions.

Returning to table 3, the results relating to the product form also demonstrate significant heterogeneity across species. Relative to the baseline fillet cut, portions, and steaks each generate both statistically significant positive and negative coefficients. It is especially surprising that portions are lower valued for the top two species of Atlantic and Sockeye, as additional labor goes into cutting fillets down to portions and additional materials go into packaging them. Rather than being a value-destroying process, portions may be a means of salvaging some value from lower-value products, similar to how other “value-added” seafood products such as breaded and battered are associated with a lower price (Roheim, Gardiner, and Asche 2007; Bronnmann and Asche 2016). For steaks, the coefficient is statistically insignificant for the species of Atlantic, Chinook, Coho, and Sockeye, but positive and significant for Pink salmon, and negative and significant for Chum. Given that the producer can choose whether to cut the fish into fillets or steaks at roughly similar costs, profit maximization should lead to their prices being similar or else the market would be all of one type. Indeed, we see this similarity in the two most frequently purchased species of Atlantic and Sockeye. Similarly, Bronnmann and Asche (2016) report that steaks have a

13. This aggregates all wild-caught salmon not from the Copper River into Alaska. Product descriptions can be used to identify 3.72% of Chinook by volume from California, and 0.01% from Washington State. Separating these out reveals a premium for California with a coefficient of 0.83 compared with 0.55 for Alaska. The Washington State coefficient of 0.54 is indistinguishable from that of Alaska. Coefficients on other characteristics change negligibly.

Table 4. Hedonic Price Function Estimates Segmented by Production Method

	Chinook		Coho	
	Wild (1)	Farmed (2)	Wild (3)	Farmed (4)
Product form				
Heads	–1.89***			
	(0.08)			
Portion	0.75***	0.34	0.58***	
	(0.22)	(0.29)	(0.13)	
Steak	–0.14**	–0.10	0.10	–0.15***
	(0.07)	(0.13)	(0.09)	(0.03)
Whole	–0.22***	–0.46***	–0.73***	
	(0.06)	(0.13)	(0.13)	
Unknown	–0.10	–0.66***	–0.05	–0.20***
	(0.07)	(0.07)	(0.08)	(0.02)
Condition				
Fresh	0.13***	1.08***	0.07	0.01
	(0.04)	(0.15)	(0.06)	(0.10)
Frozen	0.03		–0.45***	
	(0.14)		(0.03)	
Prev. frozen	–0.08		–0.12	
	(0.10)		(0.08)	
Origin				
Alaska	–0.69***		0.16*	
	(0.06)		(0.09)	
Copper Riv.			0.13***	
			(0.04)	
Chile				–0.16
				(0.11)
Constant	3.47***	1.77***	2.52***	2.32***
	(0.05)	(0.17)	(0.03)	(0.10)
Observations	19,631	4,445	24,033	3,843
R^2	0.59	0.94	0.80	0.68
SSE	689	41	781	27
Box-Cox θ	0.60	–0.07	–0.04	–0.36
χ^2 stat ($\theta = 0$)	1,541	12	11	37

Note: Two-way (county by year-month) cluster robust standard errors are in parentheses. All models employ analytic weighting on volume sold (pounds). All models include fixed effects for county, year-month, and brand name. The omitted factors are Los Angeles County, December 2016, and unbranded. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

negative but insignificant effect on price. The sign of the results for whole fish and heads and bones are in line with expectations, although it is interesting to note the difference in magnitude across species. The discount on whole fish ranges from 28% (–0.33) for Atlantic to 53% (–0.75) for Coho. The results for product form overall are not surprising, with the exception of the finding that portions receive a lower average price for Atlantic and Sockeye.

The estimated coefficients for the product condition produce some unexpected results, with both frozen and fresh products having both statistically significant positive and negative coefficients across the species. However, with the baseline being the "unknown to the researcher" condition, there is more information in the relative rankings of the observed conditions. For every species the previously frozen products have a lower price than fresh, which is to be expected given the physical changes to the meat that occur with freezing (Dawson, Al-Jeddawi, and Remington 2018). The difference between previously frozen and frozen salmon would then be that previously frozen salmon would have a shorter shelf life for the store and the household, with the only advantage of previously frozen being immediate readiness for cooking. It is likely that spoilage of the thawed products necessitates a higher margin for the grocer, explaining why the previously frozen products are more expensive than frozen. The premium for frozen Pink salmon is somewhat puzzling, but it may be tied to the abundance of fresh seasonal catch, depressing prices while the stable products can be sold out of season at a higher price. With Pink salmon representing the largest share of shelf-stable canned salmon (NMFS 2019), the relevant comparison for suppliers is likely freezing versus canning, which cannot be evaluated with these data. Overall, the results show that fresh is more valuable than previously frozen, and for most species of salmon previously frozen is more valuable than frozen.

The origin results show that there can be significant price premiums for salmon from some regions, but surprisingly Alaskan origin is not universally positive. It is worth recalling from table 1 that the origin information is missing from the dataset for 68.2% of observations that account for 84.6% of revenue. As such, it is likely that some of these unknown observations are from Alaska, Chile, Norway, et cetera,[14] which implies that these coefficients would likely be inaccurate. With that caveat, the estimated coefficients show a significant premium for the farms in Chile, Norway, and Scotland when compared with these observations with missing origin information. Within the United States, the Copper River region of Alaska is associated with premium salmon products and generates a sizable premium for Chinook and Coho. However, Alaskan origin (where known) in the data has a negative effect for Sockeye and the unknown species but a positive effect for Chinook, Chum, and Coho. The negative effects are surprising given the previous findings of price premiums for seafood produced in the same country, but the generic-origin Pacific salmon are also likely to originate from the United States. Limitations of the data prevent us from making strong statements about the impact of origin on salmon sold in California, but the price premiums on the Norwegian, Scottish, and select brands of Chilean Atlantic salmon even after accounting for brand effects suggest that building a solid regional reputation can reverse the discount on farmed salmon.

CONCLUDING REMARKS

While previous studies have sometimes evaluated "salmon" as one species, it is clear that this aggregation masks meaningful market structure differences between salmon species. Heterogeneity in consumer preferences and producer costs across species and production method, as evidenced by the significance of interaction terms, leads to less nuanced average estimates of the marginal implicit price in pooled models. This result is likely to be true not just for salmon but for many multiproduct hedonic price models in agricultural economics, including other seafood, meat, and produce.

The United States is an important but under-researched market for salmon. The market is significant internationally because the majority of the US salmon capture fishery landings are exported

14. NMFS (2019) reports year 2018 US imports of 417,822 pounds of edible fishery products imported from Chile; 189,335 pounds from Norway; and 36,649 pounds from the United Kingdom.

and a majority of the domestic consumption is imported. This article adds to the stated preference studies of salmon preferences among US consumers with the first revealed preference estimates from a hedonic price model using four years of aggregated salmon purchase data from California. Although California may not be representative of the rest of the country, it is home to approximately 12.5% of the nation's population and thus a significant market in its own right. These results are of particular use in understanding the preferences of Californians with respect to species and aquaculture, providing additional evidence of a premium for wild-caught over farmed salmon. We have also contributed to the use of retailer scanner panel data through an illustration of the data cleaning process (online appendix) and through a thorough discussion of the use of weighting in estimating hedonic price functions with aggregated purchase data.

In light of the previously consistent findings of a preference for domestically produced seafood, it is surprising that Alaskan salmon has both positive and negative willingness to pay estimates while foreign nations have positive and significant premiums. This unexpected result may be due to incomplete and possibly inaccurate origin data, but the US trade in salmon products is consistent with American consumers generally preferring Atlantic over Pacific salmon given the realized prices. However, the lack of US farmed salmon and the paucity of international Pacific salmon in the data make the ceteris paribus evaluation of origin challenging even with complete and accurate data.

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