

Lancker, Kira; Bronnmann, Julia

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# Substitution Preferences for Fish in Senegal

Kira Lancker<sup>1</sup> · Julia Bronnmann<sup>2</sup>

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## Abstract

In a marine multi-species environment, consumers' decisions may introduce interactions between species beyond biological ecosystem links. The theoretical literature shows that consumer preferences for variety can trigger a sequential (local) extinction of fish stocks. However, consumer preferences are not yet fully understood empirically, as it is uncertain how variety-loving consumers really are, in particular in specific settings such as in developing countries. In this article, we present an aggregation procedure to study consumer preferences in a highly diverse marine system. In a first step, we use co-integration analysis and aggregation theorems by Hicks and Lewbel to find groups of species that consumers find substitutable. In a second step, we use a direct quadratic almost ideal demand system (QUAIDS) to estimate price elasticities between these groups. We then quantify and compare welfare losses and spillovers from species-specific price shocks that may for example result from restoration efforts. Our case study from Senegal across 28 species reveals evidence that consumers do indeed have a preference for diversity of species on their plates.

**Keywords** Aggregation · Fish demand · Marine biodiversity · Price elasticities · QUAIDS · Substitution

**JEL Classification** Q18 · Q22 · D12 · C32 · C33

## Abbreviations

GCCT      Generalized composite commodity theorem  
LOP        Law of one price  
QUAIDS    Quadratic almost ideal demand system

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✉ Kira Lancker  
kira.lancker@idiv.de

Julia Bronnmann  
jubr@sam.sdu.dk

<sup>1</sup> Biodiversity Economics, German Center for Integrative Biodiversity Halle-Jena-Leipzig (iDiv), Puschstraße 4, 04103 Leipzig, Germany

<sup>2</sup> Department of Sociology, Environmental and Business Economics, University of Southern Denmark, Esbjerg, Denmark

# 1 Introduction

For a large share of the global population seafood plays an important role, as it builds essential nutrient complements to typical calorie-dominated diets (Pauly 2019). By contributing 17% of animal protein, and 7% of all proteins consumed globally, seafood is in particular crucial for over 3 billion people in developing countries. Due to the large variety of different species, seafood can also be a source of diversity on consumers' plates, which depends on consumer preferences. Globally, seafood is a highly traded commodity (FAO 2018), and also many developing countries benefit from seafood trade (Asche et al. 2015; Bronnmann et al. 2020). Likely, one reason for this are consumer variety preferences, as different types of fish are imperfect substitutes. All else equal, consumers' utility increases in the variety of goods (here fish) they consume. We thus interpret a lower willingness to substitute as a "preference for variety", a term that has been coined in particular in response to the Dixit-Stiglitz constant elasticity of demand function (Dixit and Stiglitz 1977).<sup>1</sup> However, one third of the fish stocks monitored by the FAO is classified as over-exploited or depleted. This may hamper the oceans' role for food provision both in terms of quantity and variety.

Theory has shown that consumer preferences induce species interactions on market level, with consequences for ecosystem resilience. If consumers readily substitute between resources, stock restoration after an adverse shock is less costly to consumers (Quaas et al. 2013). Quaas and Requate (2013) show that limited substitutability is an important determinant of stock (local) extinction under open access. In this study, we aim to add to the literature by investigating consumer preferences for different seafood species in a developing country setting with a high degree of species variety.<sup>2</sup>

Empirical analyses of fish consumer preferences mostly rely on demand system estimation, in particular in developed countries like North America, Europe and Japan (Asche et al. 1997; Bronnmann et al. 2016b; Bronnmann 2016; Wessells and Wilen 1993; Xie and Myrland 2011). But also the number of studies for developing countries is increasing (Akuffo et al. 2020; Dey 2000; Dey et al. 2011, 2005, 2008; Garcia et al. 2005). All of these are examples for direct demand systems, although there are also indirect demand systems estimated such as Barten and Bettendorf (1989) or ?.

The literature covers specific demand for the most prominent species and products. Overall, the studies conclude that the demand for fish in developed countries is inelastic. When dis-aggregated demand of specific fish species is analyzed, the demand seems to be elastic, as consumers can substitute more effectively within a highly aggregated commodity. In other words, demand is less elastic when there are limited available substitutes (Asche et al. 2007). This is also commonly observed using a multistage budgeting framework. At the second stage, the elasticities are more elastic than at the first stage (Molina 1997). Thus, elasticity estimates depend on aggregation.

For developing countries, especially for Asia, studies find high negative own price elasticities. Cross-price relationships between species groups are typically inelastic, except for high valued fish. While the majority of cross-price elasticities are positive indicating substitution relationships, they also find a number of negative relationships particularly for low

<sup>1</sup> Please note that this preference we focus on here may not include comprehensive non-use (e.g. existence) values that consumers may derive from ecological diversity.

<sup>2</sup> In this paper, we use the term 'consumer' to refer to all customers that buy fish at the ex-vessel market under study.

income groups and between processed and fresh fish types (Dey et al. 2011; Garcia et al. 2005).<sup>3</sup>

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As demand systems can only deal with a limited number of parameters, demand studies make generally use of different aggregation approaches, that follow from respective research questions. In many studies aggregation approaches are used which based on species' traits. For example, species of the same taxon are grouped. However, if consumer preferences are not based on species traits, such an aggregation approach may overlook some consumer-perceived variety. To fully understand consumer willingness to substitute, we make use of a non-trait-based approach to species aggregation that only considers market information. Furthermore, by studying a very large number of species, we prevent that potential variety is lost to the study.

In this article, we seek to answer i.) how can we study consumer preferences in a multi-species environment with a non-trait-based aggregation approach? ii.) Do consumers regard species as closer substitutes when species are biologically/ functionally similar? In other words, does it make sense to aggregate according to biological criteria? iii.) Is fish consumers' willingness to substitute limited?

To answer these questions, we use a unique species-specific catch and price panel from Senegal. The relevance of local fish supply for Senegalese food security is underlined in Lancker et al. (2019a, 2019b). Senegal is of particular interest, as trade is very limited, and variety on consumers' plates is determined by the local bio-diversity. While variety preferences are crucial in trade contexts, a case study with limited trade facilitates the analysis of variety preferences. The paper empirically evaluates revealed preferences from willingness to pay for direct use values expressed through market prices in the face of short run exogenous supply. In a first step, we use cointegration analysis and aggregation theorems by Hicks (1936) and Lewbel (1996) to aggregate 28 species into eight groups of species, where consumers find species within groups highly substitutable. Subsequently, we use a direct quadratic almost ideal demand system (QUAIDS), developed by Banks et al. (1997), to estimate substitution elasticities between these groups for the representative consumer. This approach allows us to assess to what extent consumers are willing to substitute, and to bring additional and valuable evidence for Africa to the table. We discuss policy

<sup>3</sup> A good review of seafood demand studies can be found in Asche et al. (2007).

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implications and show how our results can be used to understand resilience and the cost of stock restoration.

The article is structured as follows: Section 2 summarizes the relevant theoretical and empirical literature background. Section 3 provides an overview of the case study and the data used in this article. Section 4 presents our empirical approach. The results are presented in Sect. 5, followed by a discussion in Sect. 6 and a conclusion.

## 2 Theoretical Context: Demand-Side Interactions

Recent advances in theoretical modeling reveal the important consequences of consumer preferences in a multispecies context for ecosystem resilience (Quaas and Requate 2013; Quaas et al. 2013; Steinshamn 2017). If consumer preferences are transmitted strongly enough via price signals, and supply is somewhat elastic, fishers adapt quantities accordingly.

Consumer willingness to substitute is a determinant of species' relative equilibrium stock sizes and potential extinction<sup>5</sup>. In a multi-species system, consumers may choose between different fish species. One of the first papers to consider consumer induced species connections is Quaas and Requate (2013). The authors consider a Dixit-Stiglitz utility function (Dixit and Stiglitz 1977) of constant elasticity of substitution (CES,  $\sigma$ ) type, where species are imperfect substitutes. If consumers' willingness to substitute between species is limited ( $\sigma$  is close to unity), consumers' are said to show a "love of variety". Species are modeled as symmetric except for their natural growth rates<sup>6</sup>.

In the model, demand depends on the species own price, but also on all other prices, whereas supply depends on the stock size. For any one species, the resulting open access harvest therefore depends on the species own stock size, the other stock sizes, and on  $\sigma$ . The authors show that equilibrium stock sizes across species diverge more, the lower the elasticity of substitution ( $\sigma$ ).<sup>7</sup> Price divergence also decreases in  $\sigma$ , where the scarcer species' price is higher. As the authors note: "An intuitive explanation for this result is that stronger preferences for diversity mean that consumers have a stronger desire to consume different species in similar proportions." Then, with no or inefficient management, the least resilient species may go extinct. With falling  $\sigma$ , extinction may become sequential, and the stronger the love of variety, the more stocks may go extinct. The authors conclude that stronger preferences for diversity entail a stronger need for multi-species regulation. In line with these findings, Steinshamn (2017) reveals a conservative effect of consumer substitution on optimal steady state stocks given a sufficiently high stock elasticity in the cost function. Both models abstract from biological interactions, an assumption that we will adopt from here on as well<sup>8</sup>.

<sup>5</sup> In the analytical model, (Quaas and Requate 2013) analyze under which circumstances a stock size (of a local population) is reduced to 0, analogously to the extinction definition by ? as 100% decline over baseline abundance.

<sup>6</sup> The species are not biologically linked, and thus do not have to stem from the same ecosystem. Thus, theoretically, the result could hold for traded species from different jurisdictions. However, the model does not explicitly consider trade (costs).

<sup>7</sup> To provide more insight, we reproduce this result in Appendix A for the simplified two-species case.

<sup>8</sup> While it would be useful to study the relative importance of consumer and biological interactions, this is beyond the scope of the paper and the data-set.

Consumer willingness to substitute influences the cost of restoration after a shock. A single fish stock, that has been reduced below its optimal level, is less costly to restore if consumer demand is elastic (in terms of  $p$ -substitution), such that consumers readily buy other food (substitute). Then, the welfare loss for consumers and potentially markets higher up in the value chain from temporarily reduced harvest, e.g. through a harvesting moratorium or quota, is smaller. If, on the other hand, consumers' willingness-to-pay to continue buying fish is high (inelastic demand in terms of  $p$ -substitution), a temporary harvest reduction would incur a significant loss. For a multi-resource environment and a dynamic model, Quaas et al. (2013) show that demand side complementarity across symmetric resources may lead to multiple steady states and lower resilience of the whole system. The intuition is again that a temporary harvest reduction induces a stronger temporary reduction in instantaneous welfare.

The discussed CES type models assume for simplicity that any two species are equally good (or bad) substitutes for each other, and all commodities have an income elasticity of one (Quaas and Requate 2013). To study variety preferences in our empirical study, we opt for the demand system approach with more degrees of freedom, which has also become standard in the literature, namely the Quadratic almost ideal demand system (QUAIDS) (Banks et al. 1997).

### 3 Case Study and Data

In this article, we use data from Senegal's main fishing region Thiès Sud. Senegal is the most important fishing nation within the Canary Current large marine ecosystem. Its artisanal fishery harvests about 375 kt of fish annually (158 kt in Thiès Sud), and the national industrial fleet harvests another 43 kt (Direction des pêches maritimes 2013). Fish trade plays a minor role during the period of our data-set, such that consumers' seafood variety depends on the variety in local waters. The large number of local available species is particularly compelling, and enables us to study whether preferences for variety run along biological gradients.

Senegalese inhabitants consume 35 kg of fish per capita annually, where the consumption is as high as 50 kg along the coasts (Failler 2014; York and Gossard 2004).<sup>9</sup> The artisanal fishery mostly supplies local consumers and small-scale processors (Ba et al. 2017; Failler 2014). About 90% of the catches that reach the local markets in Thiès are either consumed fresh, traded fresh with neighbouring regions or go into local artisanal transformation. The fishery in Senegal is unregulated, except for mandatory fishing licenses purchased at a fixed price; and a range of subsidies.

We use monthly, species-specific catch data<sup>10</sup> and nominal ex-vessel prices<sup>11</sup> provided by the Senegalese maritime fisheries directorate (Direction des pêches maritimes, DPM).

<sup>9</sup> For comparison: In 2017, the consumption of fish in the European Union was 24 kg per capita (FAO 2018).

<sup>10</sup> The data does not distinguish between landings and catches, discards are assumed to be very small.

<sup>11</sup> We do not have access to consumer market prices. It must be noted that to the extent that middlemen are involved, our results may not fully reflect consumer preferences. However, in most cases consumers, or to a smaller part small-scale retailers or processors, directly purchase the fish from the fisher's family. Therefore, we assume that ex-vessel prices are a good proxy for market prices, at least for our case study.

Landings are aggregated and aggregate prices are calculated as unit prices. Prices are recorded per kg in units of the Senegalese currency Franc de la Communauté Financière d'Afrique (FCFA). Our data-set spans the period between January 2006 and May 2014, i.e. 101 months. We consider datapoints as non-missing if data for the two largest landing sites, Joal and Mbour, are available. These two landing sites account for 97–99% of the whole regional landings. The two ports lie about 32km apart from each other, a small enough distance that arbitrage by boat or small truck is established. We test whether the law of one price (LOP) holds between the two markets at species level using cointegration techniques. Our tests confirm that the two ports form one market.<sup>12</sup> If data are missing for Joal or Mbour (8 months), we use landings for the whole Thiès region (Direction des pêches maritimes 2006–2004) and apply a weighting factor. The weighting factor is calculated as a monthly mean of the Thiès Sud share in total Thiès landings on species level over the months where we do have data for both. Thus we assume that for each species and month, the relative Thiès Sud versus Thiès total landings remain relatively constant over the years. As we only have one price observation from one landing site, we have to assume that this is the overall price for the species for these datapoints. We limit our study to fish species with few missings, to avoid distortions in the time series. The remaining species account for 92% of total fish expenses from the full data-set. Furthermore, we drop eight species with extremely low budget shares far below 1%, to reduce complexity. Of the remaining observations across 28 species and 101 months, we drop 13 price outliers (< 1% of observations), that lie more than three times the interquartile range above that range, exceed three times the median price for that species and are clear visual outliers. They do not follow a pattern with respect to time step, season, or species. No outliers needed to be dropped at the low end.

Figure 1 depicts the development of the aggregate regional Thiès Sud fishery over the time horizon according to our remaining data-set across 28 species. Catches show high seasonal and inter-annual variability. The harvest fluctuates around 13 kilotons, while expenditures fluctuate around 2.4 billion FCFA. Expenditure per kg of catch shows a slight upward trend. The data shows an exceptionally low harvest and high expenditures per kg of catch in April 2007. We have neither indication of a data collection mistake, nor knowledge of a particular event that explains this exception. However, we investigate this datapoint's impact in a robustness check (see Appendix E).

Species composition concerning taxonomic family for catch and expenditures is shown in Fig. 2. Fish of the family *Clupeidae*, in particular *Sardinella spp.*, dominate catches and gain in importance over time. Since they command a low price, their nominal expenditure share is far less dominant. The second most important family on average are *Serranidae*, large, high-trophic level sea basses and groupers that command a high price. A species of particular importance in this family is *Epinephelus aeneus*, or “Thiof”, as it is called locally. This meaty white fish is part of several traditional dishes.

<sup>12</sup> We describe these techniques in detail in sect. 4.1, since they form the main methodology of our aggregation strategy. Due to brevity, we do not present results here, but they are available upon request.

## 4 Methods

### 4.1 Aggregation Approach

This section outlines our approach to aggregating numerous species into a limited number of commodity groups. Consumers make budget allocation decisions across large numbers of seafood species. However, due to large numbers of parameters to estimate, it can be difficult to estimate seafood demand systems with too many species. Therefore, in the first step, we aim to aggregate 28 species into fewer groups, where species within each group are very substitutable to the consumer, such that little variety value is lost. Aggregation should be achieved by a non-trait-based approach, using only market information.

Feasible aggregation concerns the question whether the researcher can treat a group of goods as a single commodity. Approaches to test for valid aggregation mainly follow one of two ways (Blackorby et al. 1978): The first is aggregation based on the assumption of separable preferences, imposing restrictions on the underlying utility function, following theory by Gorman (1959). Most commonly, homothetic Leontief-Sono separability is assumed for valid aggregation. However, testing for separability - typically via testing restrictions on dis-aggregated elasticities following Goldman and Uzawa (1964) - requires substantial degrees of freedom, which we do not have in our study. Testing is also subject to issues of multicollinearity (Lewbel 1996; Xie and Myrland 2011).

Hence, we employ the second approach, namely aggregation based on the generalized composite commodity theorem (GCCT) (Lewbel 1996), which generalizes the stricter Hicks-Leontief composite commodity theorem (CCT) and was previously applied to food (Reed et al. 2005) and fish (Xie and Myrland 2011) demand. Hicks-Leontief aggregation requires relative prices of aggregated commodities to remain fixed (Hicks 1936; Leontief 1936). The GCCT (Lewbel 1996) relaxes the assumption of perfect collinearity of prices within a group. Let the relative price be defined as the ratio of the price of good  $i$  to a price index, where we follow Xie and Myrland (2011) and Moschini (1995) and use a loglinear Laspeyres type price index  $P_{Index}$  according to the following function:

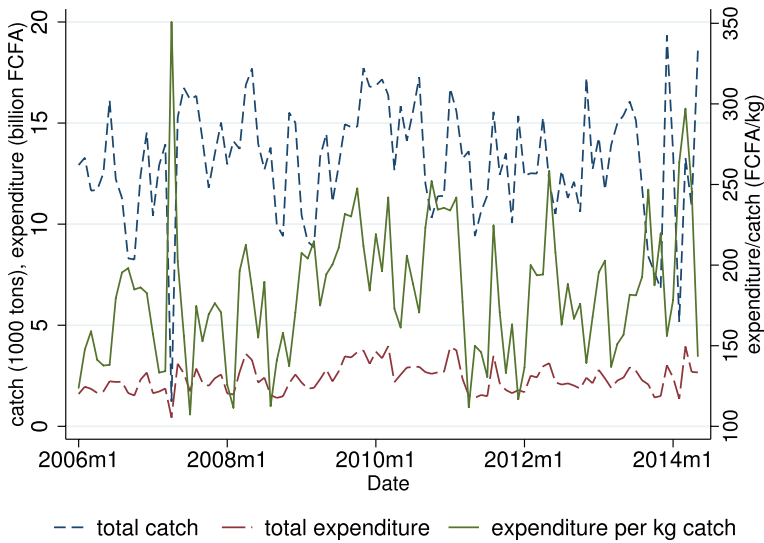
$$\log(P_{Index}) = \sum_{i=1}^n w_i^0 \log(p_{it}), \quad (1)$$

where the group consists of  $n$  species,  $p_{it}$  is species  $i$ 's price at time  $t$ , and  $w_i^0$  is its base budget share in the group, where  $t = 0$  corresponds to the first observations in our time series. This is an approximation, as we do not have knowledge of the exact underlying utility function.

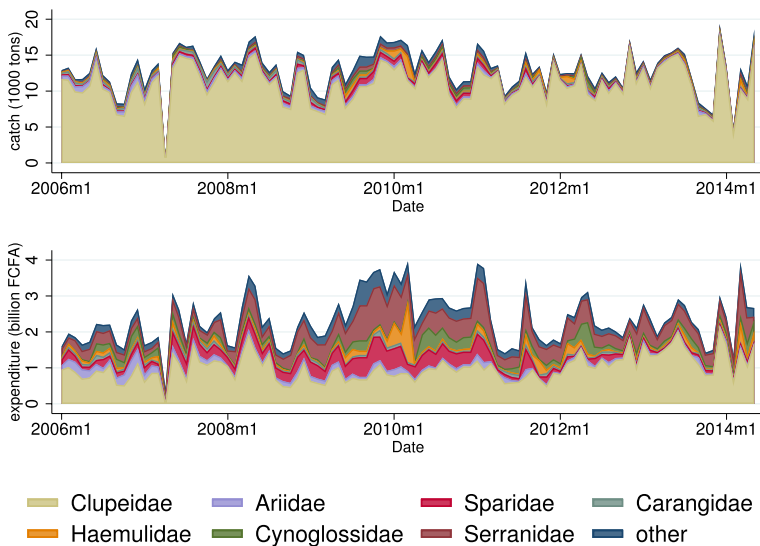
The GCCT holds if the distribution of the relative price is independent of the price index distribution. This means that one assures that, for non-stationary prices, the relative price is not co-integrated with the price index, while one tests for correlation in case of stationary prices. If the GCCT holds, a group of commodities can be treated as a single, homogeneous good in a demand system, as the remaining aggregation error is well behaved.

Under the GCCT, multiple aggregation schemes may be supported. Given 28 species, there exist over 20,000 potential groups with four species or less, let alone the larger groups. A second criterion is required to choose between alternative aggregations, ideally one that leads to an aggregation that is exhaustive across species. Previous studies use the GCCT to test a trait-based aggregation scheme. Our study aims to find such an aggregation scheme based only on market information, to preserve much of the consumer-perceived variety value. For this reason, we first rely on a stochastic version of the more stringent





**Fig. 1** Development of harvest, expenditure and nominal price level, 01.2006-05.2014



**Fig. 2** Species composition in harvest and nominal expenditures, 01.2006-05.2014

Hicks-Leontief theorem to narrow down aggregation possibilities to groupings with a high substitutability. As Asche et al. (1999) point out, the logarithmic Hicks-Leontief theorem with an error term is equivalent to the law of one price (LOP). We can therefore use the established LOP testing procedure (Asche et al. 1999). The intuition for the LOP is that there is a long-run relationship between similar commodities of integrated markets, ensured by arbitrage. For two close enough substitutes on the same market, such a relationship can

be brought about by substitution. To narrow down aggregation possibilities, we require that within a group, the LOP can not be rejected at the 5% level for all pairwise relationships.

In principle, we could instead make the LOP test even stricter. Asche et al. (1999) discuss that if all pairwise LOPs hold for a group strictly enough, the GCCT must hold. We do find that this is often, but not always the case at our level of testing (5% level). In that regard, our approach reflects a compromise between finding groups of species with a high substitution potential (as requiring the LOP test to hold at the 5% level), and allowing for enough generality to achieve sufficient aggregation by use of the more general GCCT.

To test for the LOP, we apply the standard market integration modelling based on cointegration as a complementary measure to find groups of species with high substitution value for consumers. Traditionally, the test for market integration relies on testing the relationship between two market price time series expressed in logarithms. We test the following specification of the law of one price between a pair of species  $i$  and  $j$ :

$$\ln p_{it}^1 = \zeta + \kappa \ln p_{jt}^2 + e_t \quad (2)$$

where  $p_{it}$  is the price for species  $i$  at time  $t$ . The parameter  $\zeta$  is a constant term that reflects quality differences. The error term  $e_t$  is assumed to be white noise. The parameter  $\kappa$  determines the long-run relationship between the prices. If  $\kappa=0$ , there is no relationship between the price series, while  $\kappa=1$  indicates that the LOP holds and there is complete market integration. In case of  $\kappa \neq 1$  or  $\kappa \neq 0$ , the price changes transferred from one level of the chain to the other is incomplete in the long run.

How to conduct inference for equation (2) depends on the nature of the stochastic processes of the vector of prices. In the presence of non-stationary price series, equation (2) produces spurious results. Cointegration analysis presents an alternative solution (Engle and Granger 1987). The standard unit root test based on the Augmented Dickey-Fuller model (Dickey and Fuller 1981) is used in this article.<sup>13</sup> A cointegration relationship exists when the linear combination of non-stationary variables integrated of the same order is stationary. To test for cointegration, the Johansen (1988) approach is used, which is the maximum likelihood estimator of the so-called reduced rank model.<sup>14</sup> It is based on an error correction model (ECM) representing of the Vector Auto Regressive (VAR) model with  $I(0)$  vector  $\Delta p_t$  containing log prices given by:

$$\Delta p_t = \sum_{i=1}^{k-1} \Gamma_i \Delta p_{t-i} + \Pi p_{t-1} + \mu + \varepsilon_t \quad (3)$$

The long-run relationship (i.e. the cointegration vector) is captured by the matrix  $\Pi$ , defined as  $\Pi = \zeta \kappa'$ , where  $\kappa' p_t$  is stationary if a cointegration relationship exists and the matrix  $\Pi$  has a reduced rank ( $r$ ):  $0 < r < n$ . The parameter  $\zeta$  represents the speed with which prices adjust to the equilibrium. To conclude on market integration, a rank of  $n-1$  should be identified, where  $n$  is the number of price vectors. In the following,  $\kappa$  is the parameter of interest. The LOP is tested in this setting with the restrictions  $\kappa' = [1; -1]$ .

The rank of the matrix  $\Pi$  is tested with modified chi-square tests called the trace and maximum eigenvalue statistics (Johansen, 1988). If  $\text{rank}(\Pi)=0$ , then no co-integration

<sup>13</sup> The results of the ADF tests are not presented for brevity.

<sup>14</sup> As the Johansen test uses the reduced form, the often discussed problem of endogeneity resulting from simultaneously determined prices is avoided (Asche et al. 1999).

relationship exist and if  $\text{rank}(\Pi)=n$ , then there is full rank - indicating that all the price series are stationary. Equation 2 can be written as bivariate representations:

$$\begin{pmatrix} \Delta p_{t-1}^1 \\ \Delta p_{t-1}^2 \end{pmatrix} = \begin{pmatrix} \mu_1 \\ \mu_2 \end{pmatrix} + \begin{pmatrix} \zeta_1 \\ \zeta_2 \end{pmatrix} (p_{t-1}^1 - \zeta - \kappa p_{t-1}^2) + \begin{pmatrix} \theta_{1j} & \delta_{1j} \\ \theta_{2j} & \delta_{2j} \end{pmatrix} \begin{pmatrix} \Delta p_{t-j}^1 \\ \Delta p_{t-j}^2 \end{pmatrix} + \begin{pmatrix} \varepsilon_{t-1}^1 \\ \varepsilon_{t-1}^2 \end{pmatrix} \quad (4)$$

where the parameter  $\mu$  is the deterministic component and where the number of lags,  $j = 1, 2, \dots, k$  is chosen.  $\theta$  and  $\delta$  are the own- and cross-lagged price parameters, and  $\varepsilon$  represents the error term assumed to be independent and identically distributed with zero mean and a finite variance. The lag lengths used in the estimations are selected by considering the following information criteria, where the most frequent selected lag is used: LR, FPE, AIC, SC, and HQ.<sup>15</sup> Additionally, a model search is conducted over all cases of trend assumptions and the suitable specifications chosen.

To obtain a small number of groups, we order all species according to the number of LOP connections, and start with the species with the fewest connections. We then accept the largest group possible according to the GCCT, and proceed to the remaining species with the fewest connections and so on, until all species are grouped.

## 4.2 Demand System Estimation with QUAIDS

Next, we employ the QUAIDS model, originally developed by Banks et al. (1997) to estimate price elasticities. We assume that the individual decision to consume a fish species is a result of utility maximisation subject to a budget constraint.<sup>16</sup> The model is widely used in estimating direct demand models<sup>17</sup> for seafood (e.g. Bronnmann et al. 2016a, 2019; Dey et al. 2011; Kumar et al. 2004; Garcia et al. 2005) and can be described by<sup>18</sup>:

$$w_{it} = \alpha_i + \sum_{j=1}^k \gamma_{ij} \ln(p_j) + \beta_i \ln \left[ \frac{m}{a(p)} \right] + \frac{\lambda_i}{b(p)} \left[ \ln \frac{m}{a(p)} \right]^2 + \epsilon_{it}, \quad (5)$$

where  $m$  is the total fish expenditure,  $w_{it} = \frac{p_i q_i}{m}$  is the expenditure share of fish species  $i$  in period  $t$  with  $q_i$  as the quantity of species  $i$  and  $\epsilon_{it}$  is a random disturbance. The parameters  $\alpha_i$ ,  $\beta_i$ ,  $\gamma_i$  and  $\lambda_i$  have to be estimated. The functions  $\ln a(p)$  and  $b(p)$  are translog and Cobb-Douglas price aggregator functions, respectively. The translog price aggregator is defined as  $\ln a(p) = \alpha_0 + \sum_{i=1}^k \alpha_i \ln p_i + \frac{1}{2} \sum_{i=1}^k \sum_{j=1}^k \gamma_{ij} \ln p_i \ln p_j$ , and the Cobb-Douglas price aggregator is  $b(p) = \prod_{i=1}^k p_i^{\beta_i}$ .<sup>19</sup> Both  $\ln a(p)$  and  $b(p)$  are thus specific to group  $i$ . For theoretical consistency, equation (5) is estimated under additivity

<sup>15</sup> LR: sequential modified LR test statistic (each test at the 5% level), FPE: final prediction error, AIC: Akaike information criterion, SC: Schwarz information criterion, and HQ: Hannan-Quinn information criterion.

<sup>16</sup> Representative consumer models fit aggregate consumption data reasonably well (Lewbel 1991).

<sup>17</sup> We use a direct demand system instead of the sometimes favored inverse demand system, as supply may not be sufficiently inelastic in the short term, given for example the evidence in Lancker et al. (2019b).

<sup>18</sup> For detailed derivations of expenditure share equations we refer to Banks et al. (1997).

<sup>19</sup> Following Banks et al. (1997) and Deaton and Muellbauer (1980a), we set  $\alpha_0$  to be slightly less than the lowest value of  $\ln m$  observed in the data, i.e.  $\alpha_0=12$ .

$(\sum_{i=1}^k \alpha_i = 1, \sum_{i=1}^k \beta_i = 0, \sum_{i=1}^k \gamma_{ij} = 0, \sum_{i=1}^k \lambda_i = 0)$ , homogeneity ( $\sum_{j=1}^k \gamma_{ij} = 0$ ) and symmetry constraints ( $\gamma_{ij} = \gamma_{ji}$ ).

The equations of the QUAIDS are estimated simultaneously with the nonlinear seemingly unrelated regression (NLSUR) procedure using Stata 17, which yields a consistent and asymptotically efficient parameter estimate (Deaton and Muellbauer 1980b). As imposed by the adding up restriction, the budget shares sum to unity, resulting in a singular covariance matrix of the demand system. Thus, one equation must be omitted and the parameters of the deleted equation may be retrieved using the adding up condition. We thus estimate seven equations for our eight groups. Applying the Durbin Watson test to detect auto-correlation in single equations and the Harvey LM test to test for system autocorrelation, both provided in the Stata user-written `lmanlsur` command, we detect auto-correlation in all seven equations. Thus, we estimate the system as a first order vector autoregressive model (VAR(1)), imposing a single autocorrelation parameter *AC* for all of the equations following Berndt and Savin (1975), which is compatible with the imposed adding-up condition. Visual inspection shows heteroscedasticity of residuals concerning budget shares. Single equation tests, including Pagan-Hall-tests (Pagan and Hall 1983), and system tests including Breusch-Pagan (both provided in the user-written `lmhnlstur` package) confirm the existence of heteroscedasticity at the 1% level. Therefore, we use heteroscedasticity consistent standard errors to account for this issue (White 1980). By contrast, heteroscedasticity with respect to prices or expenditure is not pronounced, except for group VII.

To calculate price and expenditure elasticities, we use the formulas provided by Banks et al. (1997). The respective own-price, cross-price and expenditure elasticities are a function of the parameter estimates of the demand system as well as the expenditure shares. In this study, we use arithmetic sample means to calculate the elasticities. Please note that by use of nominal prices, if inflation rates are positive, this implicitly adds weight to later periods.

Through the application of the QUAIDS model and given our data limitations, we adopt certain caveats which we would like to highlight briefly. First, the model regards total expenditure on fish consumption as exogenous to fish price changes. Income is assumed to change total expenditures; and our expenditure elasticity corresponds to the income elasticity (only) if the elasticity of total expenditures *m* with respect to income is unity. Second, prices are considered exogenous, even though our data consists of aggregate equilibrium quantities and prices. However, monthly supply of the upwelling-area, small-scale fishery is known to be quite variable, and driven by biomass, factor prices, seasonal and weather fluctuations. It should trace out demand reasonably well, while demand curve shifts (e.g. following systematic changes in taste) are likely more sluggish. Data scarcity and the relatively small sample size render IV estimation difficult. Tests for endogeneity (Appendix E) let us cautiously conclude to keep using our specification.

## 5 Results

### 5.1 Aggregation Results

The two combined criteria - full pairwise LOP and GCCT - lead us to define eight groups. According to our aggregation approach, consumers should find species within groups highly substitutable, indicating little preference for variety. A list of species and their group

allocation can be found in Appendix B in Table 5, along with results of the GCCT estimates (Appendix C, Table 6). LOP results and GCCT-conform grouping are visualized in Fig. 3, where each node represents a species. We numbered both species and groups in ascending order of average weighted price per kg. Connections indicate that the LOP cannot be rejected. Please note that distances between species in the plot are not meaningful, although the plotting algorithm plots nodes with similar connections closer to each other for visibility. Colors indicate group affiliation. The graph shows that most groups can be reasonably well distinguished, while others (in particular III, IV and VIII) exhibit multiple connections and are therefore difficult to distinguish.

According to the grouping results, the fish commodity landscape can be described as follows: The rightmost group I consists of cheap (80–100 FCFA/kg) small pelagic Clupeids eaten as staple food. Group II consists of cheap low level coastal predators. We find a clear distinction between the small pelagic species (group I) and medium level demersal species (groups III–V). Three types of high price commodities can be distinguished, namely the culturally important “Thiof” (group VII), a group dominated by high priced Soles and large pelagic Carangidae (group VI) and high-value large pelagic and demersal species (group VIII). Group VIII and the Soles in groups IV and VI constitute commodities also meant for export in frozen condition.

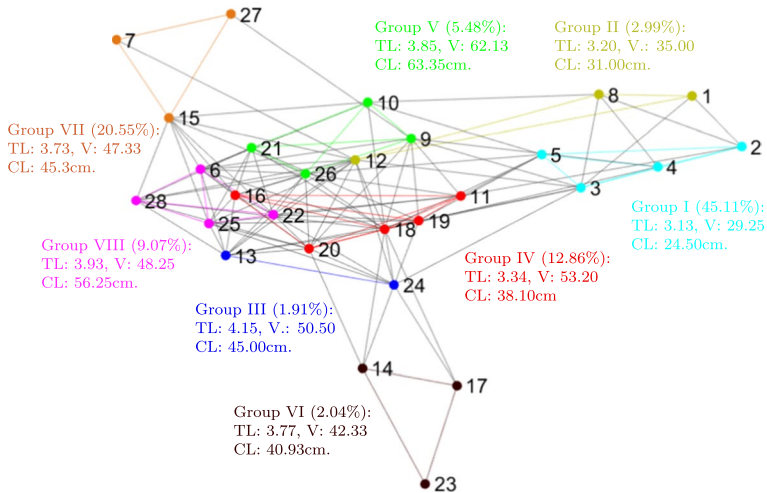
To rule out that our aggregation results stem from supply side interactions instead of consumer substitution, we explicitly consider information on gear and habitat for each group, and search for long-run relationships in harvest time series. We do not find indication that coupled or complementary harvesting affects grouping results neither in the literature nor in the harvest time series, with the exception of the littoral species in group II and *Caranx rhonchus* (#6) in group VIII.

## 5.2 QUAIDS Estimation Results

Estimation statistics are presented in Table 1, elasticity results are presented in Table 2 and parameter results in Appendix D, Table 7. The fitted model explains 77 to 98% of the variance. The Durbin-Watson statistics for each equation after correcting for autocorrelation ranges from 1.7 to 2.5, indicating that autocorrelation is adequately controlled for and the Harvey LM test no longer rejects the null hypothesis of no system autocorrelation. The autocorrelation parameter AC is 0.47 and a t-test shows that it is significant at the 1% level (see Appendix D). The residual distribution is slightly skewed left for group III and VII and skewed right for group VI. Fitting shows that for group VI and VII, this appears particularly in summer. Consequently, we check for robustness (see Appendix E) by allowing for seasonality in elasticities. Since the results remain robust, and inspecting the information criteria, we proceed without including seasonality in the chosen model.

The economic interpretation of the estimated parameters - price and expenditure effects - is best discussed in terms of elasticities, which are reported in Table 2. The approach allows us to calculate expenditure, own-price and cross-price elasticities, which we will briefly discuss.<sup>20</sup> Uncompensated elasticities take into account income and substitution effects, while compensated elasticities only account for substitution effects.

<sup>20</sup> Using the adding-up condition and Stata's nlcom command that relies on the Delta method for standard errors, we also calculate parameters and elasticities for group VIII.



**Fig. 3** Cluster of species grouped by consumer preferences. Each number belongs to a species. Higher numbers indicate a higher price. Connections indicate that the LOP holds. Colors indicate group affiliation. Share in consumer budget in brackets, along with mean trophic level (TL), vulnerability score (V) (Cheung et al. 2005) and common length (CL). For three species, no common length is available. We interpolate it by 0.6 times the maximum length

The expenditure elasticities for the eight fish groups are positive and highly significant at the 1% level, indicating increasing consumption with higher expenditure on fish. Thus, fish tends to be a normal good in Senegal. The highest expenditure elasticity is reported for group II with a value of 1.794, while the lowest expenditure elasticity pertains to group III with a value of 0.517. Our t-tests show that expenditure elasticities are significantly smaller than 1 for groups I and III and significantly larger than 1 for group VII at the 5% level. Note that this classification may change with total expenditure as our model is nonlinear in total expenditure. If our expenditure elasticity is a good indication for the income elasticity, this means that group VII is a significant luxury good, whereas groups I and III are necessities. As people become wealthier in the future, we can therefore expect that harvesting pressure increases on all groups, but proportionally stronger on group VII and weaker on groups I and III. Overall, our expenditure elasticity estimates are mostly higher than the Senegalese income elasticity for fish as aggregate commodity of 0.663 as reported by Muhammad et al. (2011) for 2005. This could be caused by the QUAIDS assumption of exogenous overall fish expenditure, given that Muhammad et al. (2011) use a multi-stage budgeting approach.

The uncompensated own-price elasticities show the expected negative signs and are statistically significant at the 5% level, indicating a negative relationship between own price and demand. The uncompensated own-price elasticities range between  $-0.463$  for group I and  $-1.379$  for group V. Thus, except for group V, demand is found to be inelastic when accounting for income and substitution effects. Our findings compare well with the finding of Muhammad et al. (2011) for aggregate food in Senegal ( $-0.75$ ), and fall in the same range as the findings by Dey et al. (2011) for several fish types in Bangladesh ( $-2.58$  to  $-0.07$ ). In general, food demand is more price elastic in developing countries (Muhammad et al. 2011).

**Table 1** QUAIDS estimation statistics (Group VIII omitted)

	eq. I	eq. II	eq. III	eq. IV	eq. V	eq. VI	eq. VII
$R^2$	.975	.841	.768	.909	.800	.808	.927
root mean sq. error	.074	.015	.012	.043	.028	.010	.059
Durbin Watson	1.8108	1.9114	2.1015	1.7546	2.4677	2.2155	1.9796
N = 100, AIC = -3207, BIC = -3077							

The focus of this study is on substitution behavior. We are interested which part of the quantity response is due to the substitution effect net of income, as summarily indicated by the compensated own-price elasticities. In Table 2, by QUAIDS restrictions, each row in the compensated price elasticities part sums to zero. This means that the compensated own-price elasticity amounts to the negative of the sum of the same group's cross-price elasticities. Compensated own-price elasticities lie between  $-0.110$  (group I) and  $-1.316$  (group V).

The compensated cross-price elasticities, which are visualized in Fig. 4, provide detailed insight into substitution connections. The identification of potential substitutes or complements is based on the sign of compensated price elasticities. We find 12 significant cross-price elasticities which are all positive and smaller than 1 in absolute terms, pointing out imperfect substitutes. However, we find for each group only a limited number of substitutes, and even in some cases no substitutes. The most substitution possibilities (3) are found for group VII. Hence, a price increase for group VII leads to an increase in the quantitative demand of group I, V and VIII. Moreover, group VII is a mutual substitute for group I, V and VIII. Consumers are willing to substitute group I, which has the highest budget share (45%), only with group VII. For group II no significant substitution effects are found.

## 6 Discussion

### 6.1 Demand Elasticity and Willingness to Substitute

Overall, we obtain mixed results concerning the willingness to substitute. Across groups, we find limited willingness to substitute, as the minimum number of groups is eight. Thus, fish is by far not a homogeneous commodity for consumers. On the other hand, within groups, our results suggest a high degree of substitutability. When considering the insights previous theoretical articles afford (see sect. 2), this means that rebuilding a stock within such a group should be facilitated by the availability of the other species within that group. This may be one contributing reason why many of the Senegalese species, even those with a high price such as *Lutjanus fulgens*, still record as “least concern” in Fishbase. More disconcerting, however, is the opposite conclusion: Overfishing of one of these species will result in a spillover of fishing pressure onto the other species within the group.

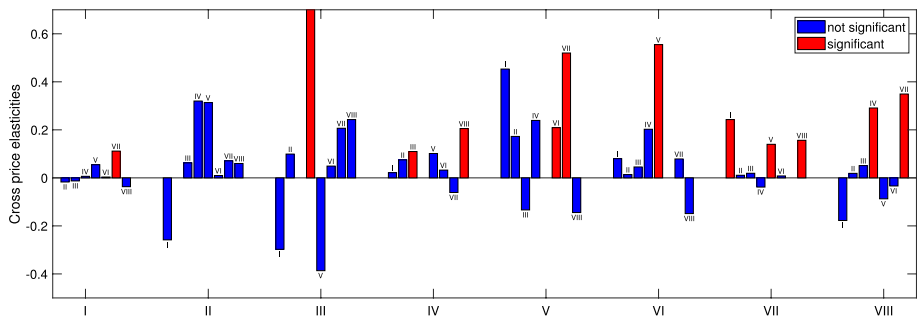
The degree of substitution between the groups can be observed from QUAIDS elasticities. This article's main interest lies in the sign and size of compensated price elasticities, which quantify the substitution effect. It allows insights about consumers' preference for variety from the ocean and the cost of restoration. To our best knowledge, this is the first study which estimates a seafood demand system for an African market.

**Table 2** Price and expenditure elasticities for fish commodity groups 2006-2014

Group I	Group II	Group III	Group IV	Group V	Group VI	Group VII	Group VIII
<b>Expenditure elasticities</b>							
0.784*** (0.089)	1.794*** (0.529)	0.517** (0.228)	1.152*** (0.102)	1.146*** (0.243)	1.002*** (0.216)	1.264*** (0.113)	1.006*** (0.167)
<b>Uncompensated price elasticities</b>							
I - <b>0.463</b> *** (0.044)	- 0.041 *** (0.014)	- 0.028** (0.011)	- 0.095** (0.044)	0.012 (0.041)	- 0.013 (0.012)	- 0.049 (0.063)	- 0.108** (0.044)
II - 1.065*** (0.281)	- <b>0.633</b> ** (0.258)	0.029 (0.135)	0.088 (0.234)	0.215 (0.217)	- 0.027 (0.081)	- 0.296* (0.178)	- 0.103 (0.205)
III - 0.530* (0.272)	0.083 (0.215)	- <b>0.671</b> ** (0.293)	0.681** (0.343)	- 0.415 (0.283)	0.038 (0.120)	0.101 (0.272)	0.196 (0.309)
IV - 0.497*** (0.139)	0.040 (0.052)	0.088* (0.050)	- <b>0.634</b> *** (0.210)	0.038 (0.097)	0.009 (0.044)	- 0.297* (0.160)	0.100 (0.111)
V - 0.063 (0.264)	0.138 (0.124)	- 0.155 (0.099)	0.091 (0.233)	- <b>1.379</b> *** (0.407)	0.185 (0.114)	0.285 (0.198)	- 0.249 (0.255)
VI - 0.371 (0.254)	- 0.017 (0.119)	0.026 (0.109)	0.073 (0.280)	0.500* (0.298)	- <b>0.847</b> *** (0.103)	- 0.126 (0.184)	- 0.240 (0.246)
VII - 0.326*** (0.124)	- 0.027 (0.031)	- 0.005 (0.026)	- 0.202* (0.110)	0.071 (0.054)	- 0.018 (0.019)	- <b>0.797</b> *** (0.142)	0.041 (0.077)
VIII - 0.630*** (0.183)	- 0.012 (0.069)	0.032 (0.065)	0.161 (0.159)	- 0.142 (0.152)	0.143 (0.168)	0.014 (0.097)	- <b>0.503</b> *** (0.188)
<b>Compensated price elasticities</b>							
I - <b>0.110</b> (0.076)	- 0.017 (0.013)	- 0.013 (0.011)	0.006 (0.041)	0.055 (0.038)	0.004 (0.011)	0.112** (0.054)	- 0.037 (0.040)
II - 0.258 (0.209)	- <b>0.579</b> ** (0.247)	0.063 (0.134)	0.320 (0.216)	0.313 (0.219)	0.010 (0.080)	0.071 (0.203)	0.060 (0.206)
III - 0.298 (0.262)	0.099 (0.212)	- <b>0.661</b> ** (0.293)	0.748** (0.344)	- 0.386 (0.284)	0.049 (0.119)	0.207 (0.282)	0.243 (0.307)
IV 0.022 (0.144)	0.075 (0.051)	0.110** (0.050)	- <b>0.484</b> ** (0.205)	0.101 (0.097)	0.032 (0.044)	- 0.061 (0.169)	0.205* (0.109)
V 0.453 (0.306)	0.173 (0.121)	- 0.133 (0.098)	0.239 (0.228)	- <b>1.316</b> *** (0.404)	0.209* (0.112)	0.520** (0.210)	- 0.144 (0.252)
VI 0.080 (0.232)	0.014 (0.118)	0.045 (0.109)	0.202 (0.276)	0.555* (0.298)	- <b>0.826</b> *** (0.104)	0.079 (0.185)	- 0.149 (0.242)
VII 0.243** (0.122)	0.011 (0.030)	0.019 (0.026)	- 0.038 (0.107)	0.140** (0.055)	0.008 (0.019)	- <b>0.539</b> *** (0.138)	0.156** (0.076)
VIII - 0.178 (0.194)	0.019 (0.068)	0.051 (0.064)	0.291* (0.156)	- 0.087 (0.151)	- 0.034 (0.055)	0.349** (0.170)	- <b>0.412</b> ** (0.184)

\* p &lt; 0.10, \*\* p &lt; 0.05, \*\*\* p &lt; 0.01. Heteroscedastic and autocorrelation consistent standard errors in parentheses





**Fig. 4** Compensated cross-price-elasticities

To provide alternative points of reference, we list results of 4 demand system studies in Table 3. We also include one study on meat commodities in South Africa, where commodity (dis-)aggregation is more straightforward.

Compensated own-price elasticities ( $-1.316$  to  $-0.110$ ) are mostly stronger than  $-0.34$  as reported by Muhammad et al. (2011) for food as an aggregate commodity. This is in line with our expectations, given the lower aggregation level in our study. It suggests that consumers substitute rather between fish groups than between food and clothes, for instance. Furthermore, we find that substitution between our fish groups is higher than between different meat types in South Africa, but way lower than between salmon and crustaceans ( $0.898$ ) (Asche et al. 1997), which was used for calibration in the theory paper by Quaas et al. (2013). This may be because our fish groups cover a wider range of commodities at the same budgeting stage, such that substitution occurs in a more evenly distributed manner in our model. It is also possible that Senegalese fish consumers, as opposed to EU consumers of salmon and crustaceans, differentiate more in the context of limited trade opportunities and a strong fishing culture, thus appreciating fish food variety. The average of our 12 significant cross-price elasticities lies at  $0.303$ , and all of them are positive. Many of our groups are, apparently, imperfect substitutes. This is in line with expectations, although a number of negative relationships were found for other developing countries (Garcia et al. 2005; Dey et al. 2011). Group I (small pelagics) seems to be a very distinct commodity from the other fish groups. Cultural characteristics may play a role: group VII is the only significant substitute for group I. Group VII contains the culturally important species *Epinephelus aeneus* (Thiof), while group I contains the prevalent *Sardinella* species, both are well-known food-fish in Senegal. Low-priced “Assorted small fish” were also found to be least substitutable by Dey et al. (2011) for Bangladesh.

Finally, our uncompensated cross-price elasticities differ substantially from the compensated ones, highlighting the relevance of income effects. For example, the income effect following an increase in the price for group I counteracts the substitution effect on group VII, such that the net effect is negative at means ( $-0.326$ ). Uncompensated cross-price effects are relevant for cross-species redistribution of harvesting pressure and of the costs of restoration. This will be discussed further in sect. 6.3.

**Table 3** Comparative results from direct demand system studies

Study	uncomp. OPE	comp. OPE	comp. CPE	region	commodities
Muhammad et al. (2011)	-0.75	-0.34		Senegal	food, aggregate comm.
Dey et al. (2011)	-2.58 to -0.07	-1.86 to -0.06	-0.28 to 1.84	Bangladesh	Carp, Tilapia, high value fish, small fish, .
Asche et al. (1997)	-3.73 to -1.56	-4.20 to -0.76	0.36 to 2.88	EU	fresh & frozen salmon, crustaceans
Taljaard et al. (2004)	-0.75 to -0.35	-0.305 to -0.161	-0.172 to 0.375	South Africa	beef, pork, chicken, mutton

OPE: own-price elasticity range, CPE: cross-price elasticity range

## 6.2 Consumer Preferences and Biological Gradients

We were also interested in whether consumer substitution is stronger when species are biologically/ functionally similar. For once, it is of interest for conservation if consumers have a preference for variety on the level of functional bio-diversity or only on species level. Moreover, the analysis offers insights concerning the underlying theory. Quaas and Requate (2013) assume fish species have different natural growth rates, but the elasticity of substitution is uniform across all stocks (for simplicity). Thus, the theory is based on the assumption that consumers substitute equally well between any two species, notwithstanding different biological traits. In line with this, we do not find evidence that consumer preferences do systematically depend on biological similarity based on information from Fishbase on size, taxonomic family, trophic level and vulnerability (Cheung et al. 2005)<sup>21</sup> (see Fig. 3 and Appendix B, Table 5). Thus, while the willingness to substitute is limited between certain species, we do not find evidence that these preferences run along functional bio-diversity gradients. The only exception are the low-trophic level small pelagics in group I, that were all grouped together. They do constitute a distinct resource. Following Hicks et al. (2019), the planktivorous feeding habit of small pelagics may be of particular importance to consumers: They contain higher concentrations of omega-3 fatty acids.

## 6.3 Management Consequences

For a manager, it is of interest to calculate welfare losses to consumers after a change in fish prices. Note that price changes will have welfare effects on consumers, fishers, inputs etc. We are here only evaluating benefits for consumers (and potentially small scale processors).<sup>22</sup> To give an example of interpretation for our QUAIDS results, we focus on the case where a whole group of species is affected. For example, a shock may affect one of the dominant species strongly, such that the potential for within-group substitution is fast exhausted, leading to price increases for the whole group. Else, if groups react homogeneously to climate impacts, such as would be expected for group I, all species in a group are affected simultaneously.

A price increase may be caused under open-access by biomass decline, when fishing costs depend negatively on stock size. Alternatively, a manager could decide to limit harvest quantities to protect a group. We have to assume that the manager has a good idea of the change in prices following such a shock.<sup>23</sup> For ease of interpretation, we will consider one group price change at a time.

Following Banks et al. (1996), we calculate the first order welfare loss (FO) of a price increase as the sum over the products of original budget shares and price changes,

<sup>21</sup> The vulnerability score (0(very low)-100 (very high)) measures a species' intrinsic vulnerability to fishing given ecological and life history characteristics. It is independent of economic pressure.

<sup>22</sup> Indeed, as discussed by Just et al. (1982) (p. 57/ 142), if fish is an essential input on consumer markets and the calculated demand functions take account of equilibrium price and quantity adjustments, welfare effects for all markets higher up in the value chain can be measured from this single market alone.

<sup>23</sup> The elicitation of absolute equilibrium welfare consequences requires a supply model, and optimally a bio-economic foundation considering species interaction. This is beyond the scope of this article, but a very important avenue for future research.

$$FO = \sum_{i=1}^8 w_i \frac{p_i^1 - p_i^0}{p_i^0} = w_i \frac{p_i^1 - p_i^0}{p_i^0}, \quad (6)$$

where the second equality results from the assumption that only a single group price changes at a time. Variable  $p_i$  is the price as before. The superscripts 0 and 1 denote variables before and after the price change. The measure gives the change in expenditures to keep the utility level at given quantities, as a percentage of sample-mean monthly total expenditures of 2.35 billion FCFA. For example,  $FO = 10\%$  amounts to 235 million FCFA (398 FCFA per capita on department level), or 2.7% of average monthly 2010 food expenditures (Agence Nationale de la Statistique et de la Démographie 2016). The  $FO$  measure gives an upper bound for the welfare loss, because consumer substitution possibilities are disregarded.

However, consumers will partly substitute the species group that has become more expensive by other commodities. To explicitly take substitution into account, our elasticity estimates are required. We calculate a second-order Taylor approximation following Banks et al. (1996) and Friedman and Levinsohn (2002). This gives the compensating variation. The compensating variation (CV, (7)) is the amount of direct income transfer that would provide consumers with the same utility level as before the change. It is the willingness to accept a price increase by consumers. CV may give a short-term measure of the willingness to pay (WTP) for species conservation if such a measure is not based on quantity restrictions, and provided the managers has knowledge of the impact of conservation on the price. If quantity restrictions apply, consumer expenditure changes have to be taken into account separately (Just et al. (1982), p. 137). In the long term, the change will have equilibrium consequences via inter-group spillovers. Thus, evaluating longer term WTP for conservation would require a dynamic model of both supply and demand.

The CV is lower, the more readily consumers substitute between groups. The second equality in (7) again results from the assumption that only a single group price changes at a time.

$$CV = FO + \frac{1}{2} \sum_{i=1}^n \sum_{j=1}^n w_i \hat{e}_{ij}^c \frac{p_i^1 - p_i^0}{p_i^0} \frac{p_j^1 - p_j^0}{p_j^0} = FO \left( 1 + \frac{1}{2} \hat{e}_{ii}^c \frac{p_i^1 - p_i^0}{p_i^0} \right) \quad (7)$$

Here,  $\hat{e}_{ij}^c$  represents our estimates for the respective compensated price elasticity between group  $i$  and  $j$ . For simplicity, we follow the literature and use the elasticities as calculated at sample means. Since all compensated own-price elasticities are negative, it follows that  $CV < FO$ . The difference between  $FO$  and  $CV$  is smaller, the less consumers are willing to substitute.

We intend to show how limited willingness to substitute influences the loss of consumer welfare. Figure 5 shows first order (FO, dashed lines) and second order (CV, solid lines) welfare losses following a single group price increase as indicated on the x-axis. The shaded areas highlight the differences between maximum loss  $FO$  and actual loss  $CV$ . The dotted red line is a line with slope 1. Now consider the role of substitution. An increase in the price of group I leads to high losses due to the large initial budget share of group I. A 10% price increase translates to a monthly first order loss of  $FO = 4.51\%$  (approximately 106 million FCFA). Due to very limited substitution possibilities, this loss is nearly fully realized.

By contrast, a welfare loss following an increase in the price of group VII is offset to a larger extent through consumer substitution. A 24% price increase for group VII would

incur a 4.93% first order loss, but this can be reduced to  $CV = 4.61\%$  by substitution, i.e. reduced by 7.52 million FCFA. This shows how over-fishing tendencies and restoration costs hinge on the degree to which consumers are willing to substitute. A major substitute are fish from group V, which inter alia contains other Serranidae and generally predatory fish similar to *Epinephelus aeneus* (Thiof).

Finally, our estimates are also of use to estimate spillover effects from singular price shocks. For example, spillovers to unmanaged groups may result from single-species (or single-group) management. It must be noted that for consumer preferences to induce species interactions as introduced in sect. 2, there needs to be strong enough price transmission, ideally symmetric price transmission.<sup>24</sup>

We use our elasticity estimates to calculate the new budget shares after a price shock. The changes in quantity demanded for an increase in the price index of group VII are reported as an example, where we assume that the prices of all species that belong to group VII increase by 24%. Then, we report the quantity change for the species with the highest baseline budget share in each group. Table 4 shows a rise in demand for groups III and V, and reduced demand for groups I, II, IV and VII. Harvesting pressure will in particular increase on group V, where the income effect is not strong enough to mitigate a strong substitution effect.

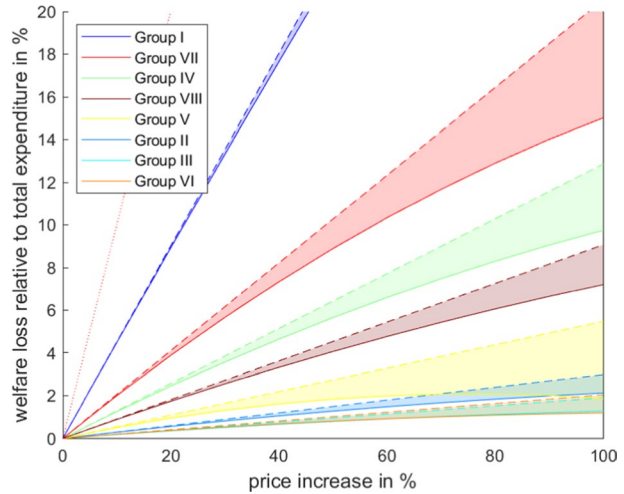
## 7 Conclusion

Overall, our study shows mixed results concerning willingness to substitute in Senegalese consumers' seafood preferences. All species have one or more close substitute, as show our aggregation results. Within groups, substitution is high. However, across groups, willingness to substitute is limited. Consumers value the variety that our oceans provide. This role of the ocean will prove difficult to recreate through aquaculture. Knowing such consumers' preferences is important to correctly gauge resilience, welfare losses and management spillovers. Using our aggregation procedure and demand system estimation, policy-makers can understand substitution patterns in their systems, e.g. to identify less resilient species. If these species are of high value to consumers, their protection is of utmost importance, but costly. For example, restoration of small pelagics is costly to consumers in Senegal exactly due to low substitution possibilities. Substitution interactions are also determinants of management spillovers. For management, a complementary issue to consider is that substitutional and complementary relationships may also occur on the supply side, for example in case of joint production, or generally where the marginal rate of substitution between inputs depends on species composition of output (Briton et al. 2021; Kirkley and Strand 1988; Squires 1987). Then, regulation of either total effort or species level output falls short of implementing optimal outcomes. The interaction between multi-product technologies and consumer induced interactions presents an interesting avenue for future research.

We look forward to future research on the topic to compare results across different ecological-economic systems. Applied multi-species modeling will benefit from a more explicit consideration of species' market interactions. Future research would be useful

<sup>24</sup> In our case study, a large part of the sales occur directly from the fisher's family to consumers, or to a smaller part to small-scale retailers or processors. Thus, ex-vessel and consumer markets are mostly the same and at least strongly linked. We can not however, test the assumption of strong price transmission, as we do not have the necessary data.

**Fig. 5** First and second order welfare loss for single price increases across groups



**Table 4** Changes in quantity demanded after a 24% increase in the price of group VII, example shows for each group the species with the highest baseline budget share

	Group I	Group II	Group III	Group IV	Group V	Group VI	Group VII	Group VIII
$q_i^{new}(t)$	7,836	212	66	82	144	9	162	27
$q_i^{old}(t)$	7,898	230	65	88	136	9	193	27
$\Delta q(\%)$	-0.79	-8.00	2.00	-7.00	6.00	0.00	-16.00	0.00

to achieve closer proximity between theoretical and empirical models, for example with respect to income elasticity effects. Prospectively, equilibrium models that incorporate species interactions are needed to support biodiversity conservation efforts. In the long term, research should incorporate substitution between fish and non-fish foods and their impact on optimal agricultural and resource policy.

## Appendix

### Outline of Quaas and Requate (2013) Result for the Two-Species Case

Quaas and Requate (2013) assume that species are symmetric except for different natural growth rates. Consumers utility depends on a numeraire and fish. Fish sub-utility is modeled as a constant elasticity of substitution utility function with elasticity  $\sigma$  across species quantities  $h_i$  and  $h_j$  for the two species  $i, j$ . Parameter  $\gamma$  describes the weight of fish consumption in utility. Prices are denoted by  $p_i$  and  $p_j$ . Marshallian demand  $h_i^{MD}$  is:

$$h_i^{MD} = \frac{\gamma p_h^{-\sigma}}{p_i^{1-\sigma} + p_j^{1-\sigma}} \quad (\text{A.8})$$

Unit harvest costs are given by  $C_i = c x_i^{-\chi}$ , where  $x_i$  is the stock size of species  $i$ ,  $c$  is a constant and  $\chi$  the stock elasticity. By  $p_i = c x_i^{-\chi}$ , the authors obtain equilibrium harvest as a function of both stock sizes and  $\sigma$ :

$$h_i = \frac{\gamma}{c} \frac{x_i^{\chi\sigma}}{x_i^{\chi(\sigma-1)} + x_j^{\chi(\sigma-1)}} \quad (\text{A.9})$$

Taking the derivative of (A.9) with respect to  $\sigma$ :

$$\frac{\partial h_i}{\partial \sigma} = h_i \chi \left[ \ln x_i - \frac{x_i^{\chi(\sigma-1)} \ln x_i + x_j^{\chi(\sigma-1)} \ln x_j}{x_i^{\chi(\sigma-1)} + x_j^{\chi(\sigma-1)}} \right] \quad (\text{A.10})$$

Note that in a stable equilibrium, the stock size of the species with higher natural growth rate will be larger than the other.

This result is used in conjunction with the differential of the equilibrium condition (logistic growth equals harvest):

$$\left( \rho_i \left( 1 - 2 \frac{x_i^{oa}}{k} \right) - \frac{d h_i^{oa}}{d x_i^{oa}} \right) \frac{d x_i^{oa}}{d \sigma} = \frac{\partial h_i}{\partial \sigma} \quad (\text{A.11})$$

The authors note that the term in brackets must be negative, as for a stable equilibrium, natural growth must increase less strongly with stock size than open-access harvest. Then,  $\frac{d x_i^{oa}}{d \sigma}$  must be positive for one stock and negative for the other.

## Species List

## GCCT Results

The GCCT holds for all eight groups. For groups I, II, V and VI, the price index is non-stationary and relative prices are stationary, providing strong evidence that the GCCT holds. For groups III and VIII, evidence for non-stationarity of the price index is weaker. For groups IV and VII, the price index is not stationary and shows low-level correlation with at least some species in the group.

Table 5 Species list

Group	Name	Family	TL	length (cm)	max. length (cm)	Vuln. (score)	mean price (FCFA/kg)	mean harvest kg/month	#
I	Ethmalosa fimbriata	Clupeidae	2.5	25	46	21	81.51	530.429	2
I	Eucinostomus melanopterus	Gerreidae	3.4	23	30	26	84.91	132.194	3
I	Sardinella maderensis	Clupeidae	3.2	25	30	33	87.00	5.709.521	4
I	Sardinella aurita	Clupeidae	3.4	25	36	37	98.62	4.930.879	5
II	Chloroscombrus chrysurus	Carangidae	3.5	25	65	28	79.61	170.639	1
II	Galeoides decadactylus	Polynemidae	3.6	30	50	31	409.76	135.200	8
II	Acanthurus monroviae	Acanthuridae	2.5	38	45	46	480.81	10.771	12
III	Pseudotolithus senegalensis	Sciaenidae	3.9	50	114	42	494.35	70.640	13
III	Cephalopholis taeniods	Serranidae	4.4	40	70	59	1929.77	3.919	24
IV	Plectorhynchus mediterraneus	Haemulidae	3.5	60	80	54	463.00	67.247	11
IV	Mugil spp	Mugilidae	2.4	37	77	41	587.49	39.989	16
IV	Diplodus spp	Sparidae	3.4	23.5	44	50.5	820.22	22.218	18
IV	Cynoglossus spp	Cynoglossidae	3.5	20	43	42.5	839.24	185.716	19
IV	Epinephelus goreensis	Serranidae	3.9	50	140	78	1057.10	79.392	20
V	Pagellus bellottii	Sparidae	3.6	25	42	42.5	428.73	150.019	9
V	Brotula barbata	Ophidiidae	3.9	50	94	52	442.60	58.231	10
V	Mycteroperca rubra	Serranidae	4.1	80	144	67	1341.09	22.348	21
V	Epinephelus caninus	Serranidae	3.8	n/a	164	87	2999.46	3.145	26
VI	Caranx crysos	Carangidae	4.1	40	70	34	516.96	15.642	14
VI	Alectis alexandrina	Carangidae	3.6	60	100	52	773.45	15.718	17
VI	Cynaptura spp	Soleidae	3.6	n/a	38	41	1602.45	17.382	23
VII	Arius spp	Ariidae	3.7	40	80	53	333.25	323.406	7
VII	Pomadourys spp	Haemulidae	3.5	36	43	37	542.00	146.976	15
VII	Epinephelus aeneus	Serranidae	4	60	120	52	3742.62	81.640	27
VIII	Caranx rhonchus	Carangidae	3.6	35	60	36	322.85	70.088	6
VIII	Pagrus caeruleostictus	Sparidae	3.7	50	90	46	1417.25	110.148	22



Table 5 (continued)

Group	Name	Family	TL	length (cm)	max. length (cm)	Vuln. (score)	mean price (FCFA/kg)	mean harvest kg/month	#
VIII	Lutjanus fulgens	Lutjanidae	4	50	60	39	1998.57	10.723	25
VIII	Epinephelus marginatus	Serranidae	4.4	n/a	150	72	4601.79	5.930	28

**Table 6** GCCT results on stationarity, incl. lags and constant, and dependence

Group	Species	Av. budget share	test stat.	corr. of rel. price with $P_I$
I	1	0.05	-5.324	
	2	0.01	-4.729	
	3	0.47	-4.480	
	4	0.47	-4.189	
	$P_I$		-3.022 <sup>+++</sup>	
II	1	0.27	-6.202	
	2	0.64	-5.648	
	3	0.10	-6.092	
	$P_I$		-2.362 <sup>+++</sup>	
III	1	0.79	-5.470	
	2	0.21	-5.470	
	$P_I$		-3.333 <sup>+</sup>	
IV	1	0.08	-4.060	0.38%
	2	0.09	-3.761 <sup>+</sup>	
	3	0.06	-5.811	-1.61%
	4	0.50	-6.466	12.77%
	5	0.25	-5.627	-23.05%
	$P_I$		-4.131	
V	1	0.51	-5.503	
	2	0.17	-4.930	
	3	0.23	-6.372	
	4	0.08	-6.058	
	$P_I$		-2.500 <sup>+++</sup>	
VI	1	0.18	-4.436	
	2	0.23	-5.753	
	3	0.60	-8.959	
	$P_I$		-2.487 <sup>+++</sup>	
VII	1	0.24	-9.022	2.69%
	2	0.13	-3.446 <sup>+</sup>	
	3	0.62	-8.956	7.55%
	$P_I$		-9.130	
VIII	1	0.13	-9.804	
	2	0.11	-7.016	
	3	0.61	-6.440	
	4	0.15	-6.374	
	$P_I$		-3.469 <sup>+</sup>	

w/o trend (G II,III,V,VII,VIII): 1% crit. value: -3.508 (+), 5% crit. value: -2.890 (++), 10% crit. value: -2.580 (+++)

w/ trend (G I,IV,VI): 1% crit. value: -4.038 (+), 5% crit. value: -3.449 (++), 10% crit. value: -3.149 (+++)

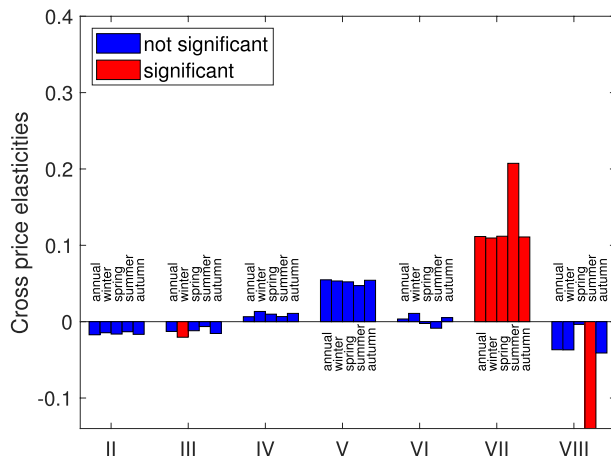
## QUAIDS Estimation Results

### Robustness

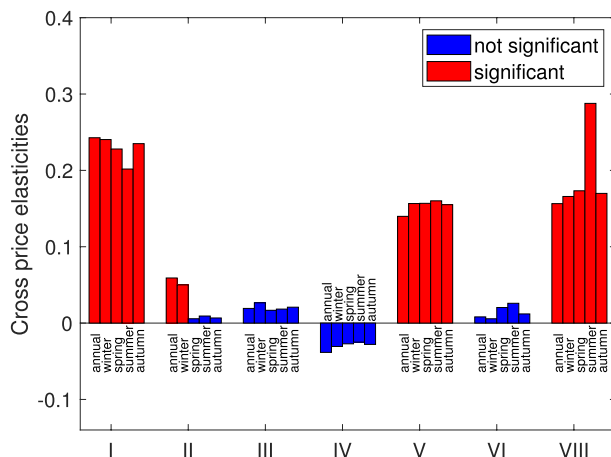
We run three robustness checks concerning the QUAIDS estimation. First, we test robustness while explicitly accounting for seasonality. Second, we look at robustness when the outlying observation in April 2007 (see sect. 3) is substituted. Third, we test for endogeneity.

First, we run our QUAIDS model with seasonality, using dummy variables for each season of winter, spring, summer and autumn. Results are very robust to including seasonality. Information criteria values AIC (-3198) and BIC (-3013) do not much improvement from including seasonality. Fitness is improved slightly for groups III, VI and VII, by 1, 1 and 2% points respectively. A likelihood ratio test, given the test statistic of 32.164 and at a difference in degrees of freedom of 21, showed that the seasonality model is not a significant improvement over our chosen model at the 5% level, although we can not reject this at the 10% level. We show as example for the full results here the changes in own- and cross-price elasticities for the traditionally important groups I (Fig. 6) and VII (Fig. 7), which are also the two largest groups with respect to mean budget share.

**Fig. 6** Seasonality of own- and cross-price elasticities for group I



**Fig. 7** Seasonality of own- and cross-price elasticities for group VII



Secondly, we observe exceptionally high prices and low harvest quantities in April 2007. While we have no indication that this is a mistake in data collection, as it has an inhomogeneous but noticeable effect on numerous species, we still investigate the impact of this observation on overall results. To this aim, we substitute the observation by the mean observed value within April observations for the respective group, and re-estimate the original QUAIDS system. Most elasticities are robust to this change. Of the significant elasticities, all but the following exceptions remain robust: The expenditure elasticity of group III becomes larger (0.517 and 0.763), while all remain significant. Furthermore, we now find the uncompensated cross-price elasticities between groups II and VII and between groups III and I to be smaller, and the former is no longer significant. More detailed results are available upon request (Table 7).

Third, there are two common types of tests for expenditure and price endogeneity. The first along the lines of Wu's variable addition test (Greene 2012) is an augmented regression approach (following e.g. Banks et al. 1997), where the residual from a first stage of regressing the potentially endogenous variable on a control is added as a variable in the actual

**Table 7** QUAIDS estimation results

Param.	Coeff./std. error		Param.	Coeff./std. error	
$\alpha_1$	-0.450***	(0.076)	$\gamma_{11}$	-0.175***	(0.036)
$\alpha_2$	0.188**	(0.082)	$\gamma_{12}$	0.034	(0.027)
$\alpha_3$	0.031	(0.041)	$\gamma_{13}$	-0.007	(0.012)
$\alpha_4$	0.124	(0.092)	$\gamma_{14}$	-0.024	(0.034)
$\alpha_5$	0.261**	(0.125)	$\gamma_{15}$	0.053	(0.033)
$\alpha_6$	0.086***	(0.029)	$\gamma_{16}$	0.012	(0.009)
$\alpha_7$	0.430**	(0.203)	$\gamma_{17}$	0.064	(0.059)
$\beta_1$	0.573***	(0.034)	$\gamma_{22}$	0.004	(0.014)
$\beta_2$	-0.083**	(0.037)	$\gamma_{23}$	0.000	(0.005)
$\beta_3$	-0.013	(0.015)	$\gamma_{24}$	0.001	(0.009)
$\beta_4$	-0.047	(0.033)	$\gamma_{25}$	0.000	(0.008)
$\beta_5$	-0.083	(0.052)	$\gamma_{26}$	-0.003	(0.003)
$\beta_6$	-0.031***	(0.012)	$\gamma_{27}$	-0.020*	(0.012)
$\beta_7$	-0.158**	(0.074)	$\gamma_{33}$	0.006	(0.006)
$\lambda_1$	-0.042***	(0.002)	$\gamma_{34}$	0.011*	(0.007)
$\lambda_2$	0.007**	(0.003)	$\gamma_{35}$	-0.009*	(0.006)
$\lambda_3$	0.000	(0.001)	$\gamma_{36}$	0.000	(0.002)
$\lambda_4$	0.004	(0.003)	$\gamma_{37}$	-0.002	(0.005)
$\lambda_5$	0.006*	(0.003)	$\gamma_{44}$	0.047	(0.029)
$\lambda_6$	0.002**	(0.001)	$\gamma_{45}$	0.002	(0.013)
$\lambda_7$	0.013**	(0.005)	$\gamma_{46}$	-0.000	(0.006)
AC	0.419***	(0.047)	$\gamma_{47}$	-0.044**	(0.021)
			$\gamma_{55}$	-0.028	(0.020)
			$\gamma_{56}$	0.007	(0.006)
			$\gamma_{57}$	0.002	(0.014)
			$\gamma_{66}$	0.002	(0.002)
			$\gamma_{67}$	-0.008*	(0.005)
			$\gamma_{77}$	0.022	(0.043)

\*  $p < 0.10$ , \*\*  $p < 0.05$ , \*\*\*  $p < 0.01$

estimation. We used annual World Bank data on GDP p.c. (World Bank 2021) and monthly dummies as controls for log expenditure and squared log expenditure. Our alternative, an index for the cost of cereal staple foods, which is available monthly, had little explanatory power as instrument. We found that the parameters in front of the residuals were not jointly significant at the 5% level (LR test with test statistic: 21.49). The same was true when using lagged values instead of monthly dummies. This test however does not work for our sample size if one also considers the seven prices as potentially endogenous, as it adds 49 (7x7) parameters to be estimated. Alternatively, the Durbin-Wu-Hausman tests for consistent parameter estimates by comparing an efficient and supposedly consistent specification (our original specification) with a consistent specification. We employed this procedure with lagged prices and the fuel price as supply shifter (from Lancker et al. 2019a) as instruments (following e.g. Brown et al. 1994). We then use the Hausman statistic to compare the performance of the consistent model, where instruments replace the potentially endogenous prices, with our chosen model. As for alternative instruments, upwelling indices as indicators of biological productivity lacked explanatory power, and data on biomass is not available. Visual inspection showed robustness of most, though not all, parameters. The Hausman test statistic is computed as  $H = (b_{orig.} - b_{inst.})'(Cov_{orig.} - Cov_{inst.})^{-1}(b_{orig.} - b_{inst.})$  for each equation, where  $b_{orig.}$  and  $Cov_{orig.}$  are the parameter vector and covariance matrix for the original specification and  $inst.$  is the subscript for the instrumented specification. Although the test statistic is supposed to be  $\chi^2$  distributed, we obtain negative Hausman statistics, which are likely attributable to the fact that the matrix in the Hausman statistic is not positive definite. According to Greene (2012), this is a known finite-sample problem, where Hausman suggests to not reject the null hypothesis of consistent estimation without controlling for endogeneity. Therefore, we cautiously conclude to use our original specification. Caveats apply: Autocorrelation in the residuals throws doubt on lagged prices to be adequate controls, but we did not identify other group-specific instruments for which data are available. Furthermore, due to our relatively small sample size, asymptotic properties of the IV regression are questionable at best.

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