Do Poor Households Pay Higher Markups in Recessions?

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Abstract

Poor and rich households greatly differ in their product choices, with the poor allocating a larger share of their spending toward cheaper goods. In recessions, all households move toward cheaper goods. Using NielsenIQ micro data, I build and calibrate a model with nonhomothetic demand and oligopolistic competition that replicates these patterns. I feed observed expenditure shifts from the Great Recession and COVID-19 into this model to isolate a demand-composition channel: demand shifts toward cheaper goods weaken cross-tier competitive pressure from premium goods, increase budget-tier markups, and tilt relative prices against baskets purchased by the poor. In the Great Recession, budget-tier markups rise 5.7 pp vs 2.3 pp for premium (relative price +2.6%); in COVID-19 they move +4.3 pp vs -2.7 pp (relative price +5.3%). Bottom money-metric welfare losses exceed symmetrically deflated spending losses by 4.9 pp (Great Recession) and 4.0 pp (COVID-19); prices account for roughly 43% of the bottom's Great Recession loss and mitigate about 10% at the top in COVID-19.

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1 Introduction

Poor and rich households systematically differ in their spending on cheap versus premium goods. Additionally, during recessions, all households shift expenditures toward more affordable goods.[1] Consequently, when a recession hits, producers of cheap versus premium goods face different changes in demand. This paper asks: How do markups respond to these unequal demand shifts in an environment in which producers' market power increases with their market share? And how does the corresponding change in relative prices affect households across the expenditure distribution? The COVID-19 episode adds salience: Survey evidence from Stantcheva (2024) documents widespread anger at price increases directed at firms, with lower-income respondents reporting sharper hardship.

I answer these questions with a multi-sector heterogeneous agent model that combines two key ingredients: nonhomothetic demand within sectors and oligopolistic competition. Within each sector, households choose among imperfectly substitutable varieties. Nonhomotheticity operates along a quality margin: the value placed on quality rises with expenditure. Higher-spending households allocate a larger share to more expensive, higher-quality varieties, while lower-spending households tilt toward cheaper options. The resulting cross-sectional patterns in purchase mixes and price elasticities match the micro evidence closely.

Sectors are oligopolistically competitive, following Atkeson and Burstein (2008). Firms' market power arises from imperfect substitutability across varieties. Small firms compete mainly on the within-sector margin, where substitution is easy and markups are low. As a firm's market share grows, competition shifts toward the across-sector margin, where substitution is harder. As a result, the firm's residual price elasticity falls and its markup rises. In short, markups are endogenously variable and increasing in a firm's market share within its sector.

[1] Bils and Klenow (1998) show luxury spending is more cyclical. Burstein, Eichenbaum, and Rebelo (2005) find lower-quality goods gain market share in recessions. Jaimovich, Rebelo, and Wong (2019) link consumption smoothing along the quality margin to stronger declines in labor demand during recessions. Jørgensen and Shen (2019) and Argente and Lee (2021) document that, facing hardship, rich and middle-class households adjust consumption at the quality margin, while the poor consumers adjust along the quantity margin.

My mechanism rests on the interaction of these two elements. In a recession, spending shifts toward more affordable, lower-quality goods within sectors. Producers of these goods gain market share and raise markups. Producers of higher-quality goods lose share and reduce markups to remain competitive. I refer to this unequal response as the *markup channel*. Quantitatively, I isolate this markup channel in two recession episodes: the Great Recession and the COVID-19 Pandemic. I feed the corresponding CEX observed distributional shifts into the pre-recession calibration while holding marginal costs and market structure fixed. In the Great Recession, markups increase more in the cheap than in the premium tier (+5.68 pp versus +2.30 pp), and the low–high relative price rises by 2.59%. In COVID-19, aggregate markups are comparatively flat but the split is sharper (+4.27 pp versus −2.71 pp), with the low-high relative price up by 5.27%. The differences across episodes turn on who enters which tier and how thick local competition is where the inflow lands—patterns the model makes precise below.

Using barcode-level NielsenIQ Homescan micro data, I document three facts about within-category quality choice and household-level price elasticities. Within narrowly defined product modules, I compute a barcode's typical shelf price per (standardized) unit and define a *premium index* that records how far that barcode's regular price sits above or below the module's typical price. Because this premium index is constructed from market-level prices it is not contaminated by household search or store choice.

I establish three facts. First, higher-expenditure households tilt toward premium varieties: bottom-quintile baskets are priced on average 0.56 standard deviations below the module mean, while top-quintile baskets are 0.42 above. Second, expenditures polarize along the premium margin. Low-expenditure households purchase mostly low-priced varieties: the interquartile range (the middle 50% of the premium index values in their baskets) runs from −0.71 to −0.34 relative to the module average. High-expenditure households purchase mostly premium varieties: their interquartile range spans +0.16 to +0.59. By contrast, middle-expenditure households mix across tiers: their median purchase is -0.05, with an interquartile range from -0.36 to $+0.45$. Third, own-price elasticities decline as a household's spending share on a variety rises. When budgets anchor low-spending households to the cheapest option, small price changes rarely reorder which good is cheapest, so there is very little within-module substitution. Similarly, higher-spending households with a strong taste for premium goods show muted responses to price changes for those varieties. The middle is more price responsive because it substitutes across quality tiers.

In the model, nonhomothetic preferences accurately reproduce the premium tilt, the inverted-U in mixing, and the elasticity gradient documented in the micro data. Mechanically, budget contractions in recessions shift purchases toward low-price corners, thin the relevant choice set, lower the demand elasticity faced by low-price producers, and raise their optimal markups. The model is set up to directly mirror the observed sectoral heterogeneity in the quality mix across product markets from NielsenIQ (the prevalence of budget versus premium suppliers) and it adopts the empirical distribution of household expenditures from the CEX. That is, both the market environment and the cross-sectional dispersion in spending are anchored directly in the data.

Parameter identification relies on a clear mapping from micro-data moments to the model's key forces. The average level of markups from PriceTrak Promo-Data disciplines the intensity of within-market competition. Markups dispersion reveals how easily consumers substitute across markets. Conditional on these forces, the price gap between premium and budget goods isolates quality-related cost differences, while, drawing on NielsenIQ's Consumer Panel, the level and dispersion of basket expensiveness (across the expenditure distribution) discipline the strength of nonhomothetic demand. Relative markups between premium and budget varieties help separate the roles of market power and tastes in explaining the premium–budget price gap. Finally, standard concentration measures from NielsenIQ's Retail Scanner Data pin down market-level demand factors that tilt baseline market shares. As validation, the calibrated model also matches untargeted cross-sectional patterns in price elasticities across quality tiers.

To quantify the markup channel, I take observed distributional shifts from the CEX and feed them into the pre-recession calibration, holding marginal costs and market structure fixed. I focus on two recession episodes: the Great Recession, with a mild decline in mean spending and a slight uptick in inequality, and COVID-19, with a sharp drop in mean spending and a compression of inequality. Two facts emerge. First, even modest aggregate price movements can mask large relative price changes along the quality margin that load disproportionately on poorer households' baskets. Second, the incidence of these changes is organized by who enters a sector (bottom versus top buyers) and how thin competition is where the inflow lands, not by sector-wide concentration.

The results are stark. In the Great Recession, markups rise more at the low tier than at the high tier (+5.68 pp versus +2.30 pp), lifting the low–high relative price by 2.59 %. In COVID-19, aggregate markups movements are comparatively flat, but the split is stronger (+4.27 pp versus $-$ 2.71 pp) and the low-high relative price jumps by 5.27 %. Decompositions show that the spending drop (in isolation) shifts demand down the quality ladder and raises low-tier pricing power, whereas higher inequality thins the elastic middle and lifts markups across all quality tiers. The compression of inequality during COVID therefore significantly counteracted the spending-drop-induced upward pressure on lower-quality markups. Tier-level competition matters precisely where the inflow lands: in sectors where the localized competition in the low-quality segment is one standard deviation above mean the relative price of these low-quality goods increases by 1.08 pp during the Great Recession. By contrast, lower competition in the high-quality tier leads to less pronounced falls in high-quality markups. A one-standard-deviation-above-mean shift of cross-sector spending in the bottom raises p_L/p_H by 2.73 pp (top lowers it by $-$ 1.63 pp). This underscores that cross-sector demand shifts are extremely consequential for markup movements.

Because baskets differ by expenditure, the markup channel generates unequal effective inflation. Using a common deflator therefore mismeasures real changes: in the Great Recession, basket prices rose by 4.32 % for the bottom quintile versus 1.79 % for the top; in COVID they rose by 3.24 % at the bottom but *fell* by -1.97 % at the top. Similar changes in symmetrically deflated spending can therefore mask vastly unequal changes in real spending. Money-metric welfare magnifies these differences: bottom-quintile welfare losses exceed symmetrically deflated spending by 4.92 pp in the Great Recession (10.68 % vs. 5.76 %) versus only 1.87 pp at the top. The markup channel alone accounts for 42.75 % of the bottom's total welfare loss in the Great Recession. By contrast, during COVID, the markup channel offsets losses at the top at a Fisher CEV of -1.83 % and mitigates about 10.55% of welfare losses among the top quintile.

Related Literature. Cementing the premise of my paper, a large body of literature studies how households adjust their consumption during recessions. Bils and Klenow (1998) show expenditures on luxuries are substantially more cyclical.[2] Burstein, Eichenbaum, and Rebelo (2005) show that lower-quality goods gain market share in recessions. Jørgensen and Shen (2019) and Argente and Lee (2021) document that rich and middle-class households smooth consumption along the quality margin, while poor consumers are more likely to adjust at the quantity margin.

Complementing my paper, a burgeoning literature studies how markups, prices, and inflation rates differ across the income distribution. Sangani (2024) documents rich households pay significantly higher retail markups for the same barcode. Here, differences in markups are due to differences in search behavior rather than product choice. Building on Kaplan and Menzio (2016), Nord (2024) connects retail price dispersion with search effort across the expenditure distribution. The response of low-quality markups to drops in spending is diametrically opposed

[2] Other seminal contributions along those lines include Browning and Crossley (2000), Aït-Sahalia, Parker, and Yogo (2004) and, more recently, Jaimovich, Rebelo, and Wong (2019) and Andreolli, Rickard, and Surico (2024).

to my predictions. As a recession hits, households throughout the economy spend less and, in turn, perceive search as less burdensome such that retailers lose market power and charge lower markups across all quality tiers. This is complementary to my argument in that I abstract away search behavior and assume perfectly competitive retailers, whereas Nord (2024) abstracts from producer competition. Kaplan and Schulhofer-Wohl (2017) and Jaravel (2019) document poor consumers experience higher inflation rates. My findings suggest these patterns are mostly driven by the sustained price impact of contractionary episodes.

Lastly, this paper contributes to the modeling of nonhomothetic demand. Building on Kongsamut, Rebelo, and Xie (2001), Buera and Kaboski (2009), Boppart (2014), and Comin, Lashkari, and Mestieri (2021), I bring nonhomotheticity inside sectors and into market power. With the recent and independent exception of Mongey and Waugh (2024), most work uses nonhomotheticity to reallocate expenditure across goods while leaving product market competition unaffected. By contrast, my framework features oligopolistically competitive firms whose strategic interactions are shaped by nonhomotheticities.

2 Model

In this section, I present a static, partial equilibrium model in which markups respond to changes in the expenditure distribution. My mechanism hinges on two critical features: nonhomothetic preferences over varieties and oligopolistic product market competition.

I first characterize pricing under a general nonhomothetic demand system and show how strategic firm interactions shape markups. I then specialize to nested nonhomothetic CES preferences that well match the micro evidence. Finally, I discuss how the implied price elasticities map into markups.

2.1 Markups under nonhomothetic demand

Firms produce varieties of differing quality and compete in oligopolistic markets. Households vary in their quality choices based on their spending, creating distinct customer compositions for producers of low- versus high-quality goods. Moreover, since households systematically differ in their price elasticities, the composition of customer base matters for market power and, therefore, markups.

Environment. There is a continuum of sectors $s \in \mathscr{S}$, each containing a finite number of quality bins $q \in \{1, \ldots, Q\}$. Within each quality bin, there are a finite number of producers $i \in \{1, \ldots, N_{qs}\}$. Each producer markets a single variety (i, q, s) and operates under constant marginal cost λ_{iqs}.

The economy is also populated by a continuum of consumers who differ in their expenditure levels y. The expenditure distribution is exogenously given and characterized by a density $g(y)$. Consumer behavior is described by general nonhomothetic Marshallian demand functions $c_{iqs}(y, \boldsymbol{p})$. Following this general discussion, I will derive a specific Marshallian demand system from a class of nonhomothetic preferences, which I later demonstrate aligns closely with the data.

Firm profits. Under Bertrand competition, firms set prices to maximize profits taking as given their competitors' prices $\boldsymbol{p}_{-iqs}$, their customers' demand functions $c_{iqs}(y, \boldsymbol{p})$, and the exogenous expenditure distribution $g(y)$. Firm profits are

$$\pi_{iqs}\big(\boldsymbol{p}, g; \lambda_{iqs}\big) = \int c_{iqs}(y, \boldsymbol{p})\big(p_{iqs} - \lambda_{iqs}\big)\, g(y)\, dy. \tag{1}$$

Note that preferences are homothetic only if $c_{iqs}(y, \boldsymbol{p})$ is linearly homogeneous in y. Consequently, with homothetic preferences, profits in (1) merely scale with aggregate expenditures, and the expenditure distribution is immaterial for the producers' profit maximization problem. By contrast, with nonhomothetic preferences, strategic firm interactions are shaped by the expenditure distribution.

Customer base. With nonhomothetic preferences, households differ in their consumption choices along the quality margin based on their spending. This variation leads producers of low- versus high-quality goods to face distinct compositions of their customer base. The consumption of (i,q,s) of a household with expenditures y relative to the aggregate consumption of (i,q,s) is denoted by

$$\tilde{c}_{iqs}(y,\boldsymbol{p},g) \equiv \frac{c_{iqs}(y,\boldsymbol{p})}{\int c_{iqs}(y,\boldsymbol{p})\, g(y)\, dy}. \tag{2}$$

This expression provides a measure of the relative importance of consumers with expenditures y for the customer base of producer (i,q,s). Note that, under homothetic preferences, the consumption of different varieties linearly scales with overall expenditure levels. That is,

$$\tilde{c}_{iqs}(y,g) = \frac{y}{\int y\, g(y)\, dy}$$

such that the customer base is homogenous across producers.

Price elasticities. With nonhomothetic preferences at the variety level, households' price elasticities for different-quality varieties also depend on spending. The price elasticity of variety (i,q,s) among consumers of type y is denoted by

$$\varepsilon_{iqs}(y,\boldsymbol{p}) \equiv \left| \frac{\partial \log c_{iqs}(y,\boldsymbol{p})}{\partial \log p_{iqs}} \right|. \tag{3}$$

With homothetic preferences we can write $c_{iqs}(y,\boldsymbol{p}) = c_{iqs}(1,\boldsymbol{p}) \cdot y$ which is tantamount to saying that price elasticities are independent of y. With nonhomothetic preferences, however, we cannot multiplicatively separate the dependence of Marshallian demand on expenditures and prices. Consequently, price elasticities differ across the expenditure distribution.

Equilibrium. The economy is represented by a marginal cost distribution $\{\lambda_{iqs}\}$, an exogenous expenditure distribution $g(y)$, and a system of Marshallian demand functions $\{(y,\boldsymbol{p}) \mapsto c_{iqs}(y,\boldsymbol{p})\}$. Firms compete in prices. The Bertrand equilibrium is then defined as a price vector $\boldsymbol{p}^* = (p^*_{iqs})$ such that consumption allocations are consistent with $\{c_{iqs}(y,\boldsymbol{p}^*)\}$ and firms' pricing decisions constitute a

Nash equilibrium. That is, $\boldsymbol{p}^*$ solves

$$\int \left(\left. \frac{\partial c_{iqs}(y, \boldsymbol{p})}{\partial p_{iqs}} \right|_{\boldsymbol{p}^*} \left(p^*_{iqs} - \lambda_{iqs}\right) + c_{iqs}(y, \boldsymbol{p}^*) \right) g(y)\, dy = 0 \quad \forall \quad (i, q, s). \tag{4}$$

Appendix B outlines an algorithm to compute this equilibrium numerically.

Markups. Firms set markups in reference to some concept of demand elasticity. With nonhomothetic demand, the relevant notion of demand elasticity for producer (i, q, s) is the cross-sectional average of consumer-specific price elasticities weighted by the corresponding relative consumption shares. That is,

$$\mathscr{E}_{iqs}(\boldsymbol{p}, g) \equiv \int \varepsilon_{iqs}(y, \boldsymbol{p})\, \tilde{c}_{iqs}(y, \boldsymbol{p}, g)\, g(y)\, dy. \tag{5}$$

Intuitively, producers consider differences in price-elasticities across the entire population, $\varepsilon_{iqs}(y, \boldsymbol{p})$, but they also take into account which consumers ultimately matter for their customer base, $\tilde{c}_{iqs}(y, \boldsymbol{p}, g)$, as well as how many of those consumers are actually present, $g(y)$.

The gross markup is defined as price over marginal cost, $\mu_{iqs} \equiv p_{iqs}/\lambda_{iqs}$. The first-order conditions in (4) dictate that profit-maximizing markups are given by the following Lerner-type formula:

$$\mu_{iqs}(\boldsymbol{p}^*, g) = \frac{\mathscr{E}_{iqs}(\boldsymbol{p}^*, g)}{\mathscr{E}_{iqs}(\boldsymbol{p}^*, g) - 1}. \tag{6}$$

The trade-offs encapsulated by this formula depend on the details of price elasticities and consumption shares, and thus on the specifics of the demand system.

2.2 Demand under variety-level nonhomotheticities

There is a continuum of consumers with nonhomothetic preferences over varieties, differing in their expenditure levels. As varieties are differentiated along a quality margin, a change in the expenditure distribution shifts spending on different-quality goods.

Preferences. Consumers choose allocations $\{c_{iqs}\}$ to maximize real consumption of a composite final good c. The aggregation of varieties into overall consumption is based on a nested nonhomothetic CES structure. At the outer nest, real consumption c is a homothetic CES aggregate of sectoral consumption c_s. Specifically, I aggregate over a continuum of sectors $\mathscr{S}$ with

$$\int_{\mathscr{S}} \left(\frac{c_s}{c}\right)^{\frac{\eta-1}{\eta}} ds = 1. \tag{7}$$

At the inner nest, nonhomotheticicies encode a quality distinction. As a result, varieties are not only imperfect substitutes but also asymmetrically differentiated along the quality margin. Sectoral consumption aggregates c_s are implicitly defined through

$$\sum_{q=1}^{Q}\sum_{i=1}^{N_{qs}} \psi_q(c_s)^{\frac{1}{\sigma}} \left(\frac{c_{iqs}}{c_s}\right)^{\frac{\sigma-1}{\sigma}} = 1 \quad \forall \quad s \in \mathscr{S} \tag{8}$$

where

$$\psi_q(c_s) \equiv \frac{\varphi_q}{c_s^{(\sigma-1)(\xi_q-1)}}$$

is a nonhomothetic taste shifter. I assume that $\sigma > \eta$ to ensure that consumption is more substitutable within sectors than across sectors. The specific functional form in (8) is based on the nonhomothetic CES preferences from Comin, Lashkari, and Mestieri (2021). The key novelty in my framework is that these preferences apply at the within-sector level. As a result, with a finite number of firms in each sector, nonhomotheticicities affect strategic firm interactions.

The parameters φ_q reflect a "consensus" on product quality, while the nonhomotheticity parameters ξ_q govern cross-sectional differences in quality appreciation. Specifically, φ_q acts as a demand shifter that is homogenous across the expenditure distribution. *Ceteris paribus*, when increasing φ_q for quality bin q, households uniformly shift spending toward varieties in this particular quality bin, irrespective of consumption levels. A key feature of nonhomothetic preferences, however, is that the appreciation of quality depends on consumption. To capture this formally, the nonhomothetic demand shifter $\psi_q(c_s)$ depends on sectoral consumption c_s. Specifically, with gross substitutes, $\psi_q(c_s)$ is strictly monotonically increasing in c_s iff $\xi_q < 1$. As a result, rich households with high consumption levels spend relatively more on low-ξ varieties, whereas poor households gravitate towards high-ξ varieties. Note that when setting $\xi_q = 1$ for all q, equations (7) and (8) specialize to the familiar homothetic nested CES structure from Atkeson and Burstein (2008).

Demand for varieties. The preferences in (7) and (8) provide markets with a demand structure. Although Marshallian demand functions are not available in closed form, the nonhomothetic CES structure allows for a great deal of characterization in terms of sharp analytical expressions.

The consumers' optimization problem is best approached in two steps. First, I focus on the consumers' within-sector expenditure minimization. In each sector s, for a given price vector $\boldsymbol{p}_s = \left(p_{iqs} : i, q\right)$, the Hicksian demand to attain aggregate sectoral consumption c_s solves

$$\min_{\{c_{iqs}\}} \left\{ \sum_{q=1}^{Q} \sum_{i=1}^{N_{qs}} p_{iqs} c_{iqs} \;\middle|\; \sum_{q=1}^{Q} \sum_{i=1}^{N_{qs}} \varphi_q^{\frac{1}{\sigma}} \left(\frac{c_{iqs}}{c_s^{\xi_q}} \right)^{\frac{\sigma-1}{\sigma}} = 1 \right\}. \tag{9}$$

The solution to this consumer program is

$$c_{iqs}(c_s, \boldsymbol{p}_s) = \psi_q(c_s) \left(\frac{p_{iqs}}{p_s(c_s, \boldsymbol{p}_s)} \right)^{-\sigma} c_s \tag{10}$$

where the nonhomothetic ideal price index is given by

$$p_s(c_s, \boldsymbol{p}_s) \equiv \left(\sum_{q=1}^{Q} \sum_{i=1}^{N_{qs}} \tilde{p}_{iqs}(c_s)^{1-\sigma} \right)^{\frac{1}{1-\sigma}} \quad \text{and} \quad \tilde{p}_{iqs}(c_s) \equiv \psi_q(c_s)^{\frac{1}{1-\sigma}}\, p_{iqs}.$$

Intuitively, we think of $\tilde{p}_{iqs}(c_s)$ as a quality-adjusted price. With nonhomothetic preferences, the appreciation of quality depends on consumption levels; therefore, the nonhomothetic ideal price index is also affected by consumption. Note that under homothetic CES preferences, with $\xi_q = \xi$ for all q, the ideal price index is homogeneous in sectoral consumption and Hicksian demand is linear in c_s.

Demand for sectoral aggregates. The quality distinction at the inner nest complicates the expenditure-minimizing choice of sectoral consumption. Consumers internalize the effect their allocations have on their nonhomothetic sectoral price indices. Taking as given the price vector $\boldsymbol{p} = \big(\boldsymbol{p}_s : s\big)$, the Hicksian demand for sectoral aggregates to attain an overall consumption of c solves

$$\inf_{\{c_s\}} \left\{ \int_{\mathscr{S}} p_s(c_s, \boldsymbol{p}_s)\, c_s \, ds \,\middle|\, \int_{\mathscr{S}} \left(\frac{c_s}{c}\right)^{\frac{\eta-1}{\eta}} ds = 1 \right\}. \tag{11}$$

Since the nonhomothetic ideal price index depends on c_s, this is akin to a homothetic expenditure-minimization problem with a non-linear pricing structure. Envisioning the continuum of sectors as a large set of cardinality S, as is common in Atkeson and Burstein (2008) settings, the corresponding first-order conditions dictate that

$$\left(\frac{c_s}{c}\right)^{\frac{\eta-1}{\eta}} = \frac{\sum_q \sum_i p_{iqs}\, c_{iqs}(c_s, \boldsymbol{p}_s)\, \xi_q}{\sum_s \sum_q \sum_i p_{iqs}\, c_{iqs}(c_s, \boldsymbol{p}_s)\, \xi_q} \quad \forall \quad s = 1, \ldots, S. \tag{12}$$

For each desired level of real consumption $c \in \mathbb{R}_+$, equations (12) are a set of S non-linear equations in S unknowns that pin down the Hicksian demand for sectoral consumption $c_s(c, \boldsymbol{p})$. Note that under homothetic CES preferences, consumers simply equate the left-hand side expression in (12) with their corresponding sectoral expenditure share.

Marshallian demand. The consumers' Marshallian demand allocates varieties $\{c_{iqs}\}$ to maximize the utility from composite real consumption c for a given budget y. Formally, the Marshallian demand functions of a consumer of type y solve

$$\arg\sup_{\{c_{iqs}\}} \left\{ c \,\middle|\, \int_{\mathscr{S}} \sum_{q=1}^{Q} \sum_{i=1}^{N_{qs}} p_{iqs} c_{iqs}\, ds \leq y \text{ and aggregators (7) and (8)} \right\}. \tag{13}$$

By duality, Hicksian demand translates into Marshallian demand. Here, "indirect" real consumption $c(y, \boldsymbol{p})$ satisfies

$$\int_{\mathscr{S}} p_s\big(c_s(c, \boldsymbol{p}), \boldsymbol{p}_s\big)\, c_s(c, \boldsymbol{p})\, ds = y. \tag{14}$$

The corresponding nonhomothetic ideal price index is then defined as

$$p(y, \boldsymbol{p}) \equiv \frac{y}{c(y, \boldsymbol{p})}. \tag{15}$$

With homothetic preferences, $c(y, \boldsymbol{p})$ is linear in y, and the ideal price index is constant across the expenditure distribution. In a slight abuse of notation, the Marshallian demand for variety (i, q, s) is henceforth denoted by

$$c_{iqs}(y, \boldsymbol{p}) = c_{iqs}\Big(c_s\big(c(y, \boldsymbol{p}), \boldsymbol{p}\big), \boldsymbol{p}_s\Big) \quad \forall \quad (i, q, s). \tag{16}$$

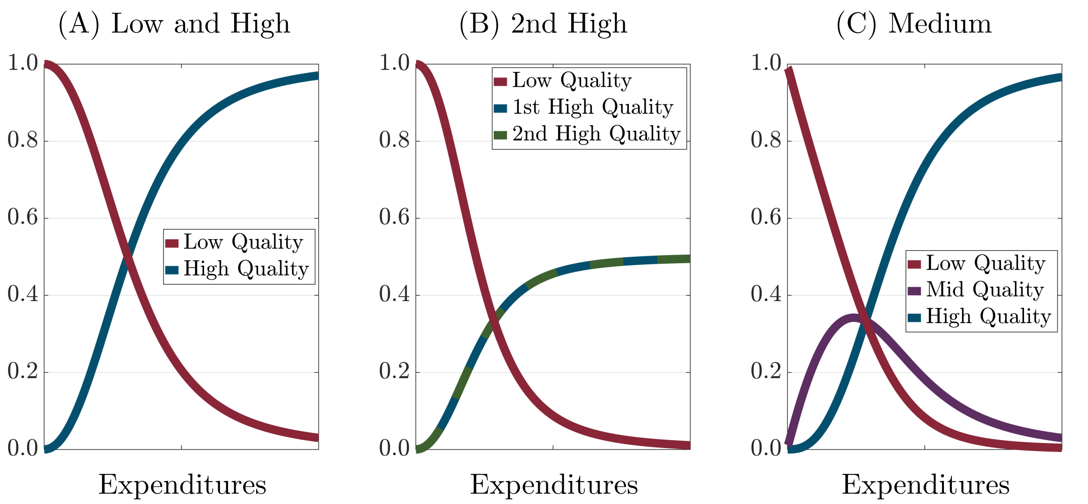
With nonhomothetic preferences, the properties of these demand functions differ across the expenditure distribution. To build intuition, the subsequent paragraphs discuss these properties.

Expenditure elasticities. Households' expenditure elasticities determine how households differ in their quality choices and, therefore, how the expenditure distribution impacts demand patterns along the quality margin. Specifically, we can examine how sectoral expenditure shares for different-ξ varieties move with sectoral consumption levels. From equation (10), Hicksian expenditure shares are given as

$$x_{iqs}(c_s, \boldsymbol{p}_s) \equiv \frac{p_{iqs}\, c_{iqs}(c_s, \boldsymbol{p}_s)}{p_s(c_s, \boldsymbol{p}_s)\, c_s} = \psi_q(c_s) \left(\frac{p_{iqs}}{p_s(c_s, \boldsymbol{p}_s)} \right)^{1-\sigma} \tag{17}$$

and depend on c_s through both the nonhomothetic demand shifter $\psi_q(c_s)$ and the nonhomothetic ideal price index $p_s(c_s, \boldsymbol{p}_s)$. We can naturally think of a variety as being of higher quality iff its expenditure share increases in sectoral real

Figure 1: Expenditure Shares as a Function of Expenditures



Panel A depicts within-sector expenditure shares for a sector with a single high-ξ (low-quality) and a single low-ξ (high-quality) variety. Panel B adds a second low-ξ (high-quality) variety, whereas panel C adds a mid-ξ (medium-quality) variety. In panel B, with multiple low-ξ varieties, rich households divide there spending among those high-quality options, while poor households continue to purchase almost exclusively low-quality goods. In panel C, medium-quality goods are predominantly consumed by middle-class households.

consumption. The elasticity of x_{iqs} with respect to c_s is given as

$$\frac{\partial \log x_{iqs}(c_s, \boldsymbol{p}_s)}{\partial \log c_s} = \Big(\sigma - 1\Big)\Big(\bar{\xi}_s(c_s, \boldsymbol{p}_s) - \xi_q\Big) \tag{18}$$

where

$$\bar{\xi}_s(c_s, \boldsymbol{p}) \equiv \sum_{q=1}^{Q} \sum_{i=1}^{N_{qs}} x_{iqs}(c_s, \boldsymbol{p}_s)\, \xi_q.$$

Note that $\bar{\xi}_s$ is the average nonhomotheticity parameter for a consumer with sectoral consumption c_s. With gross-substitutes, a particular household's expenditure share on varieties in quality bin q increases in c_s iff ξ_q is below this household's average nonhomotheticity parameter. It follows that the lowest-ξ variety is unambiguously perceived as being of high quality and vice versa. Since $\lim_{c_s \to 0} \bar{\xi}_s(c_s, \boldsymbol{p}) = \max\left\{\xi_q\right\}$, poor consumers tend to perceive mid-ξ varieties as being of high quality, while richer household, for whom $\bar{\xi}_s \to \min\left\{\xi_q\right\}$, view the exact same varieties as inferior. Generally, for $Q > 2$ and $q \notin \partial\mathcal{Q}$, quality is not an intrinsic feature of a variety but rather a matter of perception, which is contingent on consumption and, therefore, ultimately expenditure levels.

Figure 1 illustrates expenditure shares as a function of expenditures. Panel A shows these Engel curves for a sector with $Q = 2$ and $N_{qs} = 1$. As spending increases, consumers allocate a greater portion of their budget to the high-quality (low-ξ) variety. Panel B introduces a second high-quality option. Here, poor consumers' spending remains concentrated on the low-quality (high-ξ) option, while rich households divide their spending between the two high-quality varieties. Panel C, introduces a mid-ξ variety. As poor consumers spend more, they allocate a larger share of their spending to this mid-ξ variety, which they perceive as high quality. Conversely, more affluent consumers decrease their relative spending on what they now view as an inferior product.

Price elasticities. Households' price elasticities play a crucial role in determining how demand responds to price changes and are thus directly linked to firms' market power. With nonhomothetic preferences, these elasticities differ across households based on their expenditure levels.

Under Bertrand competition, the price elasticity of variety (i, q, s) among consumers with expenditures y is given as

$$\varepsilon_{iqs}(y, \boldsymbol{p}) = \Big(1 - x_{iqs}(y, \boldsymbol{p})\Big)\, \sigma + x_{iqs}(y, \boldsymbol{p})\ \eta\ \zeta_{qs}(y, \boldsymbol{p}) \tag{19}$$

where

$$\zeta_{qs}(y, \boldsymbol{p}) \equiv \frac{\Big(\sigma\, \bar{\xi}_s(y, \boldsymbol{p}) + (1-\sigma)\, \xi_q\Big)^2}{\sigma\, \eta\ \bar{\xi}_s(y, \boldsymbol{p})^2 + (1-\sigma)\, \eta\ \bar{\xi}^2_s(y, \boldsymbol{p}) + (1-\eta)\, \bar{\xi}_s(y, \boldsymbol{p})}$$

and

$$\bar{\xi}^2_s(c_s, \boldsymbol{p}) \equiv \sum_{q=1}^{Q} \sum_{i=1}^{N_{qs}} x_{iqs}(y, \boldsymbol{p})\, \xi^2_q.$$

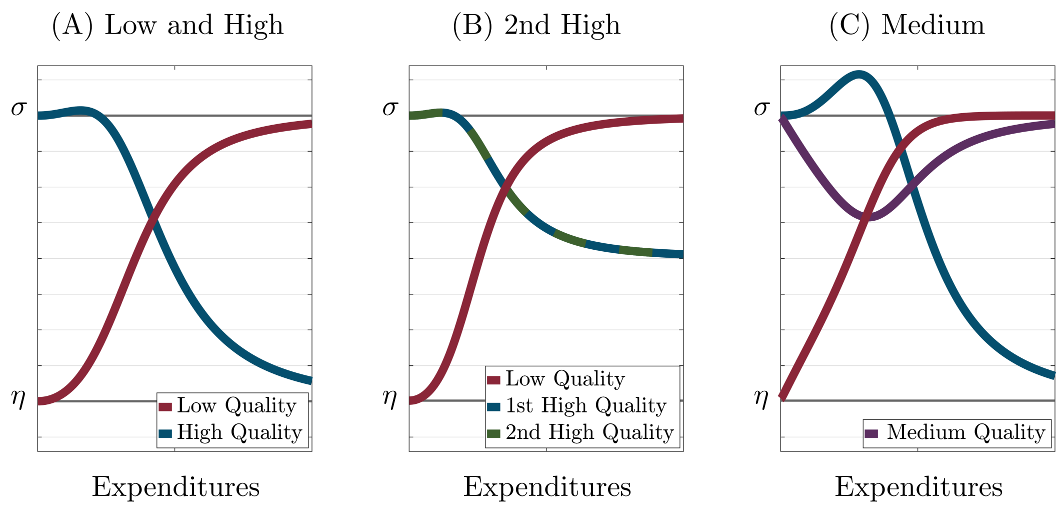
The price elasticity in equation (19) represents a convex combination of the within-sector elasticity of substitution (σ) and a modified form of the across-sector elasticity of substitution ($\eta \times \zeta_{qs}$). The additional term $\zeta_{qs}(y, \boldsymbol{p})$ captures that prices do not only influence sectoral price indices but, through their impact on c_s, also affect how households perceive quality. Thus, $\zeta_{qs}(y, \boldsymbol{p})$ is a reflection of the "non-linearity" in $p_s(c_s, \boldsymbol{p}_s)$. Note that ζ_{qs} depends on $(y, \boldsymbol{p})$ only through expenditure shares. Consequently, cross-sectional heterogeneity in price elasticities is fully explained by x_{iqs}.

The key insight from equation (19) is that the larger the expenditure share a particular household allocates to a specific variety (i, q, s), the less price-elastic they are regarding that variety. For instance, in sectors with a single inexpensive low-quality variety, consumers with $y \to 0$ allocate almost 100% of their spending to this option. Since they do not view pricier high-quality varieties within the same sector as feasible substitutes, their price elasticity approaches η. From the perspective of these consumers, there is effectively no within-sector competition, but only across-sector competition with other low-quality varieties in different sectors. This highlights the importance of competition within quality bins: when poor households can distribute their spending over multiple low-quality options, their weight on η decreases.

Similarly, rich consumers gravitate toward pricier, higher-quality products without significant regard for price. In markets with a single high-quality option, they do not even consider substitution for lower-quality alternatives. In sectors with multiple high-quality option, however, they recognize their substitutability and are thus more responsive when the price of one of those goods changes. More interestingly, middle-class households, which consume a mixture of low- and high-quality goods, are, in principle, willing to substitute along the quality margin and, therefore, comparatively price-elastic in either direction. This holds true even in relatively concentrated sectors.

Figure 2 illustrates price elasticities as a function of expenditure. Panel A shows price elasticities for $Q = 1$ and $N_{qs} = 1$. Panel B depicts a less concentrated sector with two high-quality options. Here, poor consumers remain price-inelastic for the low-quality option, as the additional pricier high-quality variety does not qualify as an affordable substitute. By contrast, with more options in the high-quality

Figure 2: Price Elasticities as a Function of Expenditures



Panel A depicts price-elasticities for a sector with a single high-ξ (low-quality) and a single low-ξ (high-quality) variety. Panel B adds a second low-ξ (high-quality) variety, whereas panel C adds a mid-ξ (medium-quality) variety. In panel A, poor and rich households, whose consumption is concentrated, are price-inelastic regarding their preferred ξ. Middle-class households are price-elastic in either direction. Adding a second high-quality variety in panel B means that rich consumers have more options and are thus more price-elastic vis-à-vis high-quality goods. These options are not feasible for poor households and their price-elasticities remain unchanged. In panel C, the addition of a medium-quality good increases price-elasticities for middle-class households. Poor and rich households remain price-inelastic for the bulk of their consumption.

segment, richer consumers become more price-responsive relative to panel A. Panel C introduces an additional medium-quality variety, which is primarily consumed by middle-class households. This additional option increases middle-class price elasticities across the board.

Note that the price elasticity in (19) generalizes a more familiar setting. When $\xi_q = 1$ for all q, preferences specialize to a nested homothetic CES structure. In that case, expenditure shares are constant across the expenditure distribution, and $\zeta_{qs}(y, \boldsymbol{p})$ is identically equal to one. Consequently, the price elasticity in (19) collapses into the standard expression from Atkeson and Burstein (2008)

$$\varepsilon_{iqs}(\boldsymbol{p}) = \Big(1 - x_{iqs}(\boldsymbol{p})\Big)\,\sigma + x_{iqs}(\boldsymbol{p})\,\eta.$$

Moreover, if there is no distinction between within- and across-sector substitutability ($\sigma = \eta$) we eliminate the effects of granularity and the price elasticity even further specializes to the expression obtained under monopolistic competition, $\varepsilon_{iqs} = \sigma$.

Markups. The demand elasticity $\mathscr{E}_{iqs}(\boldsymbol{p}, g)$ in equation (5) is given as the cross-sectionally averaged consumer-specific price elasticity ε_{iqs} weighted by relative consumption shares $\tilde{c}_{iqs}$. From Figure 1, we have seen that consumers in the tails of $g(y)$ concentrate their spending on either low-quality or high-quality goods, which makes them relatively price inelastic for most of their purchases. By contrast, middle-class households consume a mixture of low-, medium-, and high-ξ varieties. With this greater willingness to substitute along the quality margin, they have effectively more options and are therefore comparatively price elastic.

When setting markups, firms trade off the loss of business from relatively price-elastic middle-class customers against the rents they could extract from their less elastic customer segments. Consider, say, a producer selling an expensive, high-quality variety. Their customer base consists of price-elastic middle-class households and extremely inelastic consumers in the right tail of $g(y)$. When considering a price increase, this producer weighs the loss of business from middle-class consumers against the higher margin earned from their price-insensitive affluent customer segment. This trade-off accounts for the mass of each consumer type within the population. Strategic price-setting, therefore, depends on the expenditure distribution.

3 Micro evidence for nonhomothetic demand

This section tests three empirical implications of the demand system: (i) higher-expenditure households tilt toward premium varieties; (ii) consumption is polarized along the premium margin (tails specialize, middle mixes); and (iii) household-specific price elasticities are lower for varieties that command larger spending shares. I construct barcode-level measures of expensiveness and map each fact to a model counterpart.

Data. The main source is the *NielsenIQ Homescan Panel* (Chicago Booth Kilts Center), an unbalanced panel of roughly 50,000 U.S. households, 2004-2022, recording barcode-level quantities and prices of fast-moving consumer goods purchased for personal use across many retail outlets. The panel includes rich demographics and is designed to be projectable to the entire U.S. economy. Homescan covers roughly 30–40% of goods expenditure and about 15% of total expenditure. I complement this with *NielsenIQ Retail Scanner* weekly store-level prices and volumes from about 35,000–50,000 grocery and drug stores (over 90 chains), covering over half of U.S. sales volume, and with wholesale prices from *PriceTrak PromoData* (12 wholesalers, 2006–2012) to study markups relative to barcode expensiveness.

Premium score. Barcodes are grouped into narrowly defined product modules m (e.g., fresh apples, mozzarella, instant coffee) of close substitutes. To ensure within-module comparability of price data, I convert barcode quantities and prices to module-specific base units; if no natural conversion exists (e.g., count vs. ounces), I segment modules accordingly. For region r and year t, the barcode's regular price per base unit is

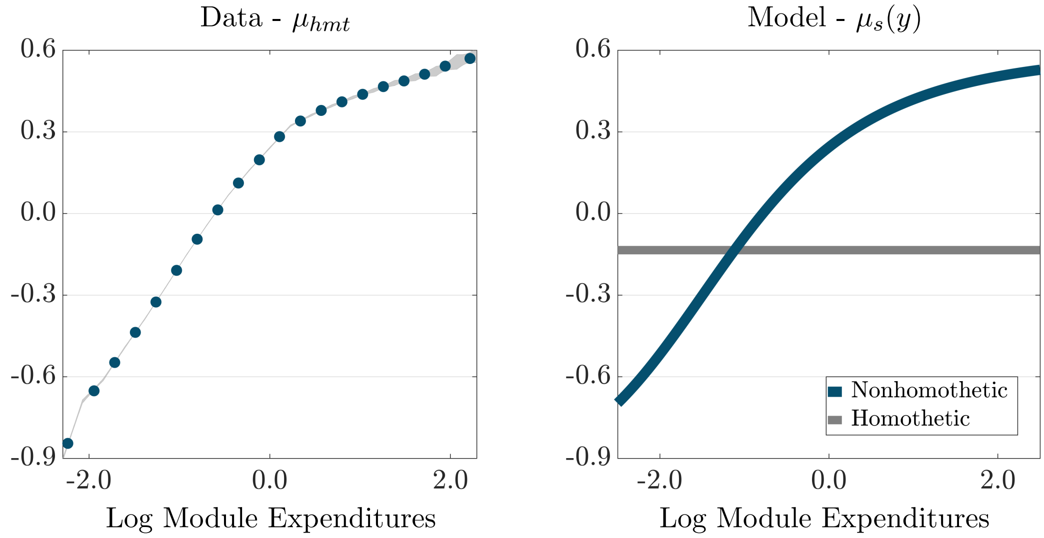
$$\text{price}_{irt} \equiv \frac{\sum_{h\in r} \text{expenditure}_{iht}}{\sum_{h\in r} \text{quantity}_{iht}},$$

a market-level object averaging across households and stores. Within a module, higher price_{irt} means a more expensive variety. To compare across modules and purge regional/time heterogeneity, I residualize price_{irt} on module, region, module×region, and year fixed effects and normalize by a module-clustered residual dispersion:

$$\text{premium}_{irt} \equiv \frac{\text{price}_{irt} - \alpha_{\text{module}} - \alpha_{\text{region}} - \alpha_{\text{module}\times\text{region}} - \alpha_{\text{year}}}{\sigma_{\text{module}}}.$$

Hence, premium_{irt} is the number of within-module standard deviations by which barcode i is priced above the typical level in (r, t).

Figure 3: Taste for Premium Goods Increases with Expenditures



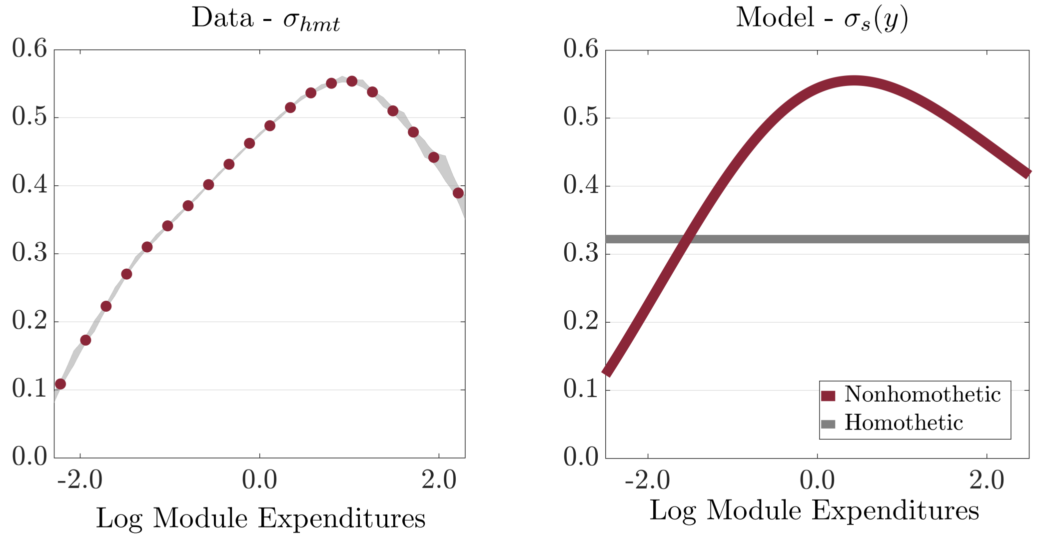
The left-hand side panel depicts a binscatter of the logarithm of region/time normalized household module-expenditures on the x-axis against household premium-indices μ_{hmt} on the y-axis. The construction of confidence band for this binscatter follows Cattaneo *et al.* (2023). The right-hand side panel plots the model-implied relationship between expenditures and household premium indices under homothetic as well as non-homothetic preferences.

Fact 1: Rich households spend more on premium goods. For household h, module m, and year t, define the quantity-weighted premium index

$$\mu_{hmt} = \sum_{i \in m} \frac{\text{quantity}_{iht}}{\sum_{j \in m} \text{quantity}_{jht}} \text{premium}_{ir(h)t}.$$

Note that μ_{hmt} depends on household behavior only through quantities; the price component in μ_{hmt} is at the market-level and does not reflect search behavior or store characteristics. Figure 3 is a binscatter of μ_{hmt} against (region–time normalized) log module expenditure and shows a strong premium tilt: low-spending households purchase varieties about priced about 0.9 standard deviations below the module norm, and the premium index rises monotonically with expenditure. High-spending households do not simply purchase larger quantities of the same varieties, but rather more expensive varieties priced about 0.6 standard deviations

Figure 4: Polarization in Consumption Patterns



The left-hand side panel depicts a binscatter of the logarithm of region/time normalized household module-expenditures on the x-axis against household premium-dispersion σ_{hmt} on the y-axis. The construction of confidence band for this binscatter follows Cattaneo *et al.* (2023). The right-hand side panel plots the model-implied relationship between expenditures and household premium dispersion under homothetic as well as nonhomothetic preferences.

above what is typical for the module. In the model, the counterpart is

$$\mu_s(y) = \sum_{q=1}^{Q} \sum_{i=1}^{N} \frac{c_{iqs}(y, \mathbf{p})}{\sum_{q,i} c_{iqs}(y, \mathbf{p})} p_{iqs}.$$

The right-hand side panel of Figure 3 shows that homothetic preferences imply constant quantity weights and no tilt, whereas nonhomothetic preferences accurately generate the observed positive slope.

Fact 2: Consumption polarization. I define the within-household dispersion in the expensiveness of purchased goods as

$$\sigma^2_{hmt} = \sum_{i \in m} \frac{\text{quantity}_{iht}}{\sum_{j \in m} \text{quantity}_{jht}} \left(\text{premium}_{ir(h)t} - \mu_{hmt}\right)^2.$$

Small σ_{hmt} indicates specialization at one end of the premium margin; large values indicate mixing. Empirically, σ_{hmt} is low for both low- and high-expenditure households and highest in the middle. Households in the bottom tertile of the module expenditure distribution, mix less than those in the middle and at the top, by a factor of 2.73 and 1.96, respectively. Poor households almost exclusively opt for inexpensive goods, whereas wealthier households predominantly consume premium goods. By contrast, middle-class households purchase a broad mixture of varieties along the premium margin. The left-hand side panel of Figure 4 illustrates the corresponding inverted U-shaped as a binscatter of expenditures against households premium dispersion. Once again, the model analogue,

$$\sigma_s^2(y) = \sum_{q=1}^{Q} \sum_{i=1}^{N} \frac{c_{iqs}(y, \mathbf{p})}{\sum_{q,i} c_{iqs}(y, \mathbf{p})} \left(p_{iqs} - \mu_s(y)\right)^2,$$

varies only with basket composition. While homothetic preferences cannot produce the inverted U, the right-hand side panel of Figure 4 shows that nonhomothetic preferences do.

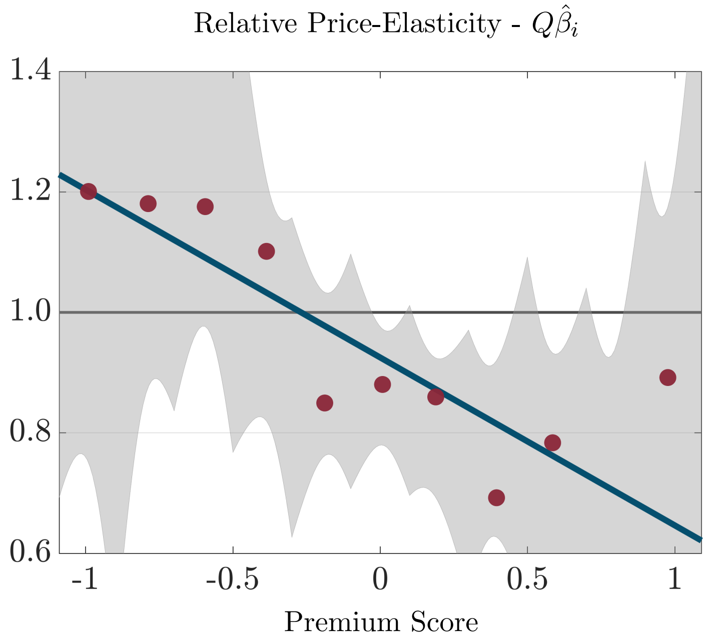
Fact 3: Households' price-elasticities decline in spending shares. I stratify households within each (m, t) by their premium index: *premium consumers* are in the upper tertile of $h \mapsto \mu_{hmt}$ and *basic consumers* are in the lower tertile. For barcode i, I estimate log-linear demand for each group using observations (i, h, t) with h in that group for $m(i)$:

$$\begin{aligned} \text{log quantity}_{iht} = \; & \alpha_{ih}^{\text{prm}} + \alpha_{ir}^{\text{prm}} + \alpha_{it}^{\text{prm}} + \beta_i^{\text{prm}} \text{log price}_{iht} \\ & + \sum\nolimits_{j \in \mathscr{K}_{iht}} \beta_{ij}^{\text{prm}} \text{log price}_{jr(h)t} + \gamma_i^{\text{prm}} \text{ log expenditure}_{ht} + \epsilon_{iht}^{\text{prm}}. \end{aligned} \tag{20}$$

$\mathscr{K}_{iht}$ includes barcodes in $m(i)$ available to h given observed store visits (constructed from Homescan trips and Retail Scanner prices). β_i^{prm} is the own-price elasticity among premium consumers; I run the symmetric regression for basic consumers to obtain $\hat{\beta}_i^{\text{bsc}}$ and define the relative elasticity

$$Q\hat{\beta}_i \equiv \hat{\beta}_i^{\text{bsc}} / \hat{\beta}_i^{\text{prm}}.$$

Figure 5: Differential Price Elasticities Across the Expenditure Distribution



This graph depicts a binscatter of the barcode-level premium scores on the x-axis against the corresponding relative price elasticities $Q\hat{\beta}_i$ for rich vs poor households. Confidence bands are constructed following Cattaneo, Crump, Farrell, and Feng (2023). For details see Section 3.

Endogeneity of price$_{iht}$ is addressed via a Hausman-style instrument: the average price of barcode i in year t excluding region $r(h)$. Identification leverages within household/region/time price variation. After controlling for module fixed effects, a binscatter of $Q\hat{\beta}_i$ against barcode premium shows that for inexpensive varieties, basic consumers are less price-elastic than premium consumers. The reverse holds for premium varieties. Figure 5 illustrates.

The regularity here is that as households concentrate their spending on goods within a particular price range, they effectively encounter fewer options and are, consequently, less price-elastic. Intuitively, poor consumers, who routinely buy the least expensive options, show minimal substitution responses to minor price changes. Similarly, wealthy consumers, with a strong appetite for premium goods, exhibit very little consumption response to price fluctuations for these pricier varieties. As a result, household-level price elasticities decline in spending shares.[3]

[3] Appendix A estimates barcode-level elasticities as a function of within-module expenditure

Table 1: Targets, Fit, and Validation

Moment	Source	Data	Model
Internally Calibrated			
Relative price (high/low)	NielsenIQ HMS & RMS	1.23	1.228
Premium index (mid/poor)	NielsenIQ HMS	1.10	1.082
Premium index (rich/poor)	NielsenIQ HMS	1.18	1.195
Premium dispersion (mid/poor)	NielsenIQ HMS	2.62	2.304
Premium dispersion (rich/poor)	NielsenIQ HMS	1.85	1.699
Aggregate markup	PriceTrak PromoData	1.36	1.355
Dispersion markups	PriceTrak PromoData	0.12	0.118
Relative markup (high/low)	PriceTrak PromoData	1.02	1.002
Average HHI	NielsenIQ RMS & GS1	0.15	0.146
Dispersion HHI	NielsenIQ RMS & GS1	0.08	0.079
Untargeted			
Relative Elasticity (low)	NielsenIQ HMS & RMS	1.18	1.268
Relative Elasticity (high)	NielsenIQ HMS & RMS	0.76	0.697

Quality distinction based on k-means clustering of premium scores. Households spending classification (poor/mid/rich) corresponds to tertiles of the within-module expenditure (after winsorization). Markup data from retail markups computed as $\text{markup}_{irt} = \text{price}_{irt}/\text{wholesale-cost}_{irt}$ using Retail Scanner prices and wholesale cost from PriceTrak PromoData. GS1 maps barcodes into firms and delivers module HHIs.

4 Quantification

In this section I outline my calibration strategy. I parameterize my nonhomothetic demand structure using moments from the NielsenIQ Homescan Consumer Panel and use CEX expenditures to capture the distribution of household spending beyond fast-moving consumer goods.

Quantitative model and sector composition. For tractability, I make a binary quality distinction with $q \in \{\text{low}, \text{high}\}$. Firms' marginal costs λ_q are quality-dependent, with no differences within tiers. Sectors are fully characterized by the number of firms operating in each quality tier. To align sector compositions with the data, within each NielsenIQ module I classify products into cheap versus expensive using k-means clustering ($k = 2$) on normalized barcode-level prices. In each module cell I then count $\left(N_{\text{low}}, N_{\text{high}}\right)$ and group cells into "sector

shares and shows the same pattern.

Table 2: Calibrated Parameters and Identification

Parameter	Value	Targets	Identification
$\lambda_{\text{low}}, \lambda_{\text{high}}$	1.00, 1.22	Relative price Aggregate markup Relative markup	Relative price pins $\lambda_{\text{high}}/\lambda_{\text{low}}$; levels close on the relative and aggregate markup.
$\xi_{\text{high}}, \xi_{\text{low}}$	1.00, 0.71	Premium indices Premium dispersion Relative markup	Nonhomothetic tilt across the expenditure distribution; relative markup further limits markup wedges vs. cost differences.
$\varphi_{\text{low}}, \varphi_{\text{high}}$	0.95, 1.59	Average HHI Dispersion HHI	Market share shifts to match the mean and dispersion of concentration.
σ	17.95	Aggregate markup	Greater substitutability within sectors reduces mean markups uniformly across constellations.
η	0.36	Markup dispersion	Greater substitutability across sectors compresses the cross-constellation spread in markups.
ν	26,327.36	Premium indices Premium dispersion	Scales CEX expenditures so basket price indices and dispersion align across groups.

Parameter values and brief heuristic identification argument. Since $\eta \neq 1$ the scale of λ and ξ is technically identified, but identification is weak within the relevant region of the parameter space. Due to nonhomotheticities λ and φ are separately identified. ν maps symmetrically-deflated USD in the expenditure distribution into units that are meaningful given ξ. The model is overidentified.

constellations." I retain the 21 most frequent constellations, which together cover about 85% of markets in the data; each constellation's measure equals its empirical frequency. The expenditure distribution is given by the empirical distribution of baseline CEX data in normal times.

Identification & calibration. Since tougher intra-tier competition lowers mean markups, the level of markups identifies the within-sector elasticity σ. The dispersion of markups across sector constellations disciplines the across-sector elasticity η: greater substitutability across sectors raises sectoral demand elasticities and compresses the cross-constellation spread in markups. Conditional on (σ, η), the relative price of high versus low quality pins $\lambda_{\text{high}}/\lambda_{\text{low}}$, and the markup level determines their levels. The premium indices and dispersion are direct reflections of nonhomothetic tastes across the expenditure distribution and, therefore, identify the nonhomotheticity parameters $\{\xi_q\}$ jointly with the expenditure scale ν. The relative markup provides an additional restriction that limits the contribution of markup wedges to the high/low price gap, helping separate nonhomothetic tastes from cost differences. Finally, the average and dispersion of concentration (HHI) identify the bin-specific demand shifters $\{\varphi_q\}$, which tilt baseline market shares holding technology and elasticities fixed.

Left untargeted, I compare price-elasticity patterns by quality and across the expenditure distribution; the model reproduces higher elasticities for low-quality varieties among poorer households and the converse for richer households.

5 Quantifying the markup channel

I show how recession-induced changes in the expenditure distribution reshape markups and, in turn, relative prices. I study the Great Recession and COVID-19, two episodes in which spending tilted toward lower-quality varieties. Calibrated to pre-episode data, the model maps the observed shifts into markup and price changes. In both episodes, the implied price changes disproportionately burden low-income households, and the predicted relative-price movements match the data.

5.1 Expenditure shifts in the data

I track CEX non-committed nondurable expenditures[4] (symmetrically PCE-deflated) across two episodes: the Great Recession and COVID-19. Mean spending falls in both, more during COVID (-16.15%) than in the Great Recession (-5.33%). Dispersion moves in opposite directions: the Great Recession slightly widens inequality (Gini $+1.20$ pp), whereas COVID compresses it (Gini -1.40 pp). The standard deviation also declines, especially during COVID (-19.31%). Table 3 reports pre/post moments and changes.

[4]To focus on readily reallocated spending, I exclude, among others, housing payments (B001-B006), education (D105, D108), insurance (A102, A107), vehicle (D001), and furniture (D002) purchases.

Table 3: Expenditure Moments: Great Recession and COVID-19

	Mean Spending	Standard Deviation	Gini	P80/P20
Great Recession				
Pre	12,932.51	7,868.67	0.312	2.837
Post	12,242.58	7,778.51	0.324	3.013
Δ	$-5.33\,\%$	$-1.15\,\%$	$+1.20\,\text{pp}$	$+0.176$
COVID-19 Pandemic				
Pre	12,836.33	9,305.10	0.355	2.924
Post	10,763.20	7,508.22	0.341	2.658
Δ	$-16.15\,\%$	$-19.31\,\%$	$-1.40\,\text{pp}$	-0.266

Moments computed on CEX non-committed nondurable expenditures, symmetrically PCE-deflated. "Δ" reports percentage changes for mean and standard deviation; level changes in percentage points (pp) for Gini and level changes for P80/P20.

5.2 An unequal markup response

I feed the observed shifts in the expenditure distribution into the pre-episode calibration, holding marginal costs and market structure fixed. The model delivers a pronouncedly unequal markup response across quality tiers: low-quality markups rise in both episodes, while high-quality markups increase slightly in the Great Recession but fall during COVID-19. With costs fixed, these markup movements map directly into prices. Table 4 reports the full set of markup and price changes.

Great Recession. The aggregate markup rises by +3.10 pp. This is driven by a sharp increase for low-quality goods (+5.68 pp) and a smaller increase for high-quality goods (+2.30 pp). The implied price changes are +4.50% (low) and +1.81% (high), yielding only a modest rise in the low-to-high relative price of +2.59%.

COVID-19. The aggregate markup is essentially flat (−0.13 pp) because a sizable increase for low quality (+4.27 pp) is offset by a decline for high quality (−2.71 pp). Prices move accordingly: +3.27% (low) versus −1.98% (high). The aggregate price index shifts little (+0.16%), but the relative price of low to high

Table 4: Unequal Markup and Price Responses by Quality

	Great Recession		**COVID-19 Pandemic**	
	$\Delta\mu$	Δp	$\Delta\mu$	Δp
Aggregate	$+\,3.10$ pp	$+\,2.45\,\%$	$-\,0.13$ pp	$+\,0.16\,\%$
Low quality	$+\,5.68$ pp	$+\,4.50\,\%$	$+\,4.27$ pp	$+\,3.27\,\%$
High quality	$+\,2.30$ pp	$+\,1.81\,\%$	$-\,2.71$ pp	$-\,1.98\,\%$
Relative price (low/high)	$+\,2.59\,\%$		$+\,5.27\,\%$	

$\Delta\mu$ reports changes in markups in percentage points. Δp reports the associated price changes with marginal costs held fixed. Price aggregates are expenditure-weighted. The bottom row reports the percentage change in the relative price, defined as the low- to high-quality price ratio p_L/p_H.

quality climbs substantially (+5.27%). The steeper increase in the low-relative-to-high price during COVID-19 implies a disproportionate burden on poorer households, consistent with evidence on the salience of price increases among lower-income consumers (Stantcheva, 2024).

5.3 Inspecting the mechanism

I decompose the markup response with two within-model counterfactuals per episode, shifting the expenditure distribution to isolate the markup effects of a drop in spending and, separately, a change in inequality.

Mean-only (inequality fixed). Let μ_0, G_0 be pre-episode and μ_1, G_1 post-episode moments. To isolate the mean change,

$$y^{(M)} = \frac{\mu_1}{\mu_0}\, y_0,$$

which preserves the Lorenz curve and, therefore, measures of inequality (Gini, P80/P20).

Table 5: Counterfactuals: Mean-Only versus Dispersion-Only

	Great Recession		**COVID-19**	
	Mean-only	Inequality-only	Mean-only	Inequality-only
Markups				
Aggregate	-0.03 pp	$+0.84$ pp	-1.31 pp	-0.98 pp
Low quality	$+0.75$ pp	$+0.79$ pp	$+1.80$ pp	-0.84 pp
High quality	-0.92 pp	$+0.92$ pp	-3.28 pp	-1.04 pp
Relative price (low/high)				
Change	$+1.26$ %	$+0.11$ %	$+3.84$ %	$+0.14$ %

Columns report counterfactual changes holding marginal costs and market structure at pre-episode values. "Mean-only" rescales all expenditures by μ_1/μ_0 (inequality fixed). "Inequality-only" dilates expenditures to match G_1 at fixed mean μ_0. Price aggregates are expenditure-weighted.

Inequality-only (mean fixed). To isolate the inequality change, apply a mean-preserving dilation that matches G_1:

$$y^{(D)} = \frac{G_1}{G_0}\,(y_0 - \mu_0) + \mu_0.$$

Findings. Mean-only shifts raise low-quality markups and lower high-quality markups in both episodes and explain most of the relative-price change (see Table 5). In the Great Recession, markups move by $+0.75$ pp (low) and -0.92 pp (high), pushing the low/high relative price up 1.26%. During COVID-19 the pattern strengthens ($+1.80$ pp low and -3.28 pp high) with a 3.84% rise in the relative price. Lower mean spending tilts demand toward cheaper varieties, expanding their residual market power while weaker top demand disciplines high-quality markups.

Dispersion-only shifts move markups with the sign of the inequality change and contribute little to relative prices. When inequality rises in the Great Recession, markups increase across tiers ($+0.79$ pp low and $+0.92$ pp high) with a small relative-price effect ($+0.11$%). When inequality falls during COVID-19, markups decline across tiers (-0.84 pp and -1.04 pp) and the relative-price effect remains modest (+0.14%). A thinner middle reduces the weight of elastic cross-tier shoppers and softens competitive pressure; a thicker middle does the opposite.

5.4 Local competition and cross-sector demand shifts

Conventional concentration has little predictive power for relative-price or markup movements. Recessionary shifts in the expenditure distribution reallocate demand across sectors and toward lower price (quality) segments and markups move most where competition is thin at the segment that absorbs demand.

Localized competition. I measure segment-specific rivalry by kernel-weighting rivals in price space. With pairwise similarity weights $w_{ij}^s = K\big(h^{-1}|p_{is} - p_{js}|\big)$ and pre-episode shares $\boldsymbol{x}_s$, I define a generalized sectoral Kernel-Herfindahl–Hirschman Index as $\text{K-HHI}_s = \boldsymbol{x}_s' W_s \boldsymbol{x}_s$ which downweights rivals at far-away prices. Note that this nests the conventional HHI as $h \to 0$. To profile concentration at a given price tier, I define local shares

$$\tilde{x}_{is}(p) = \frac{x_{is} K\big(h^{-1}\,|p - p_{is}|\big)}{\sum_i x_{is} K\big(h^{-1}\,|p - p_{is}|\big)}, \qquad \text{K-HHI}_s(p) = \sum_i \tilde{x}_{is}(p)^2, \tag{21}$$

which I evaluate at $p \in \{p_s^L, p_s^H\}$ in my two-tier baseline. Intuitively, $\text{K-HHI}_s(p_s^L)$ summarizes the effective concentration among low-price rivals actually facing the inframarginal buyers that expand in a recession.

Within-model regressions. To quantify how much of the model-implied variation in relative prices and markup changes is attributable to demand reallocation versus segment-specific concentration, I regress $\Delta \log(p_s^L/p_s^H)$ and $\Delta \log \mu_{is}$ on independent variables including HHI_s, K-HHI_s, the localized $\text{K-HHI}_s(p)$, and cross-sector spending shifts for below- and above-median households ($\Delta x_s^{\text{bottom}}$, Δx_s^{top}). All covariates are z-scored, so coefficients read as standardized semi-elasticities.

Relative prices. As reported in Table 6, conventional HHIs are almost completely uninformative for relative price changes ($R^2 = 0.025$ during the Great Recession). Before accounting for demand shifts, sectors with one standard deviation higher concentration in the low-quality segment saw a +1.08 pp larger change in the low-high relative price, whereas those with higher high-segment concentration saw it lower by −1.01 pp . Intuitively, as spending flows into the low tier, sectors with less competition in this tier better translate inflows into pricing power; conversely, with lower competition in the high tier, outflows do not force strong downward adjustments in p_s^H. Once cross-sector spending shifts along the expenditure distribution are added, demand composition is first order: an inflow of bottom spending raises the relative price of lower-quality goods by +2.73 pp per shift standard deviation. By contrast, an inflow of top spending increases pricing power among high-quality producers and thereby lowers the low-high relative price by −1.63 pp . The fit is near-exhaustive ($R^2 \approx 0.997$), and localized K-HHIs add no incremental power. This is consistent with concentration mattering only in the segment that absorbs the inflow.

COVID exhibits the same pattern, even more demand-driven: the conventional HHI has only a modest association with relative price changes. After adding bottom/top shifts and initial exposures, even localized K-HHIs become immaterial. Bottom-spending inflows raise the low-high ratio by +2.62 log points and top spending lowers it by −1.59. The independent variables again explain virtually all variation ($R^2 \approx 0.986$).

Markup changes. Table 7 shows that, for within-model variety changes in markups, HHI again carries little information. By contrast, quality is highly informative: low-quality varieties' markups rose by +2.49 pp more than high-quality in the Great Recession and by +5.22 in COVID. Localized competition matters modestly absent demand controls but becomes negligible once demand shifts are included. In the saturated specification, incidence follows who enters and what they buy. During the Great Recession, a one–standard-deviation higher inflow in bottom spending lowers high-quality markups by −0.85 pp and raises low-quality markups by +2.83 (a +3.71 differential). A one–standard-deviation higher inflow in top spending, by contrast, raises high-quality by +2.43 pp and leaves low-quality essentially unchanged at −0.01 after a −2.44 interaction. In COVID, bottom inflows are near zero for high quality at (+0.03) pp and add +2.20 for low quality; top inflows add +1.27 to high and +0.19 to low. These controls account for virtually all variation ($R^2 = 0.992$ and 0.997), indicating that buyer reallocation and tier-specific pass-through organize markup changes.

5.5 Welfare across the expenditure distribution

Recessions tilt relative prices across quality tiers. Because baskets are nonhomothetic, households face different effective inflation, so identical nominal spending drops need not translate into equal real (let alone welfare) losses. The markup channel raises low-quality relative prices, loading inflation onto necessity-heavy baskets and biasing common deflators toward understating losses at the bottom and, at times, overstating them at the top.

Real expenditures by quintile. Under homothetic CES a single index applies to all households. With nonhomothetic preferences, the relevant index is basket-specific. I therefore deflate each CEX quintile using its model-implied price index evaluated at pre-episode shares. Table 8 reports: (1) symmetrically deflated spending, (2) the model-implied basket price change from the markup channel, and (3) implied real spending. Column 2 is exactly the error from using a common deflator.

In the Great Recession, basket prices rose by 4.32% for the bottom quintile and by 1.79% for the top. A common deflator thus understates the real contraction by 4.32 pp at the bottom (moving from −10.08% to −5.76%) and by only 1.79 pp at the top (−4.72% vs. −2.93%). This pattern is even sharper in COVID-19, where prices increased by 3.24% for low-end baskets but fell by −1.97% for high-end baskets.

Table 8: Nominal and Real Spending by Expenditure Quintile

	Δ Nominal Spending	Δ Price Index	Δ Real Spending
Great Recession			
Q1 (lowest)	– 5.76 %	4.32 %	– 10.08 %
Q2	– 9.60 %	3.01 %	– 12.60 %
Q3	– 9.32 %	1.69 %	– 11.01 %
Q4	– 4.09 %	1.85 %	– 5.94 %
Q5 (highest)	– 2.93 %	1.79 %	– 4.72 %
COVID-19 Pandemic			
Q1 (lowest)	– 7.82 %	3.24 %	– 11.06 %
Q2	– 13.51 %	2.44 %	– 15.95 %
Q3	– 17.76 %	– 0.38 %	– 17.39 %
Q4	– 16.28 %	– 1.63 %	– 14.65 %
Q5 (highest)	– 17.61 %	– 1.97 %	– 15.64 %

Spending changes (column 1) are CEX expenditures deflated by the symmetric PCE index. Column 2 reports model-implied basket-specific price changes from recession markups. Column 3 equals column 1 minus column 2 (log approximation).

From real expenditures to welfare. Even basket-deflated spending is not a welfare measure under nonhomotheticity, because the ideal Konüs index between price vectors $\boldsymbol{p}_0$ and $\boldsymbol{p}_1$ depends on reference utility u:

$$\mathbb{P}(u \,|\, \boldsymbol{p}_1, \boldsymbol{p}_0) = \frac{e(\boldsymbol{p}_1, u)}{e(\boldsymbol{p}_0, u)}. \tag{22}$$

where $e(\boldsymbol{p}, u)$ denote the expenditure function dual to the Marshallian demand problem from Equation 13. As a result, equal changes in deflated spending can mask unequal welfare changes. I therefore quantify total welfare changes over the recession period through money-metric utility, i.e., a consumption-equivalent variation $\mathrm{cv}_h^{\mathrm{total}}$ satisfying

$$v(y_{0h}, \boldsymbol{p}_0) = v(y_{1h}\,\mathrm{cv}_h^{\mathrm{total}}, \boldsymbol{p}_1), \qquad \mathrm{cv}_h^{\mathrm{total}} = \frac{e(\boldsymbol{p}_1, u_{0h})}{y_{1h}}, \tag{23}$$

with $u_{0h} = v(y_{0h}, \boldsymbol{p}_0)$. Note that, because v is ordinal, $\mathrm{cv}_h^{\mathrm{total}}$ is invariant to its cardinalization.

To isolate the welfare consequences of prices movements, I hold spending fixed and compute Laspeyres and Paasche CEVs,

$$\mathrm{cv}_h^L = \mathbb{P}(u_{0h} \mid \boldsymbol{p}_1, \boldsymbol{p}_0), \qquad \mathrm{cv}_h^P = \mathbb{P}(u_{1h} \mid \boldsymbol{p}_1, \boldsymbol{p}_0), \tag{24}$$

and summarize with the superlative Fisher, $\mathrm{cv}_h^F = \sqrt{\mathrm{cv}_h^L \mathrm{cv}_h^P}$. In Table 9, I report the price contribution to the total welfare change as

$$\mathrm{PriceShare}_h = \frac{\ln \mathrm{cv}_h^F}{\ln \mathrm{cv}_h^{\mathrm{total}}}, \tag{25}$$

where negative values indicate that prices fell and mitigated losses.

Welfare. Money-metric welfare magnifies distributional gaps. For the bottom quintile, welfare losses exceed symmetrically deflated spending losses by 4.92 pp in the Great Recession (10.68% vs. 5.76%) and by 3.96 pp in COVID-19 (11.78% vs. 7.82%). Unfavorable price movements alone require a 4.28% compensation in the Great Recession and account for 42.75% of the bottom quintile's total welfare loss. Intuitively, among those poorer households necessity-heavy baskets leave little scope to cut quantities, so prices do most of the harm. At the top, the welfare-spending wedge is modest (1.87 pp in the Great Recession and 1.46 pp in COVID), and in COVID prices actually cushion losses: the Fisher CEV implies a willingness to pay of about 1.83% of income to preserve the favorable price change. That is, the price response mitigates roughly 10.55% of total welfare losses among the rich. The top welfare-spending gap is mostly due to quality downgrading, a channel that is clearly secondary to the price burden borne by the poor.

Table 9: Welfare by Expenditure Quintile

	Welfare Cost	Price Share	Laspeyres	Paasche	Fisher
Great Recession					
Q1 (lowest)	10.68 %	42.75 %	4.26 %	4.29 %	4.28 %
Q2	14.33 %	27.29 %	3.60 %	3.83 %	3.71 %
Q3	13.26 %	22.73 %	2.65 %	2.97 %	2.81 %
Q4	6.26 %	32.24 %	1.90 %	1.97 %	1.94 %
Q5 (highest)	4.80 %	36.67 %	1.73 %	1.74 %	1.74 %
COVID-19 Pandemic					
Q1 (lowest)	11.78 %	27.08 %	3.03 %	3.08 %	3.05 %
Q2	15.53 %	17.05 %	2.14 %	2.55 %	2.34 %
Q3	21.96 %	3.82 %	0.25 %	1.30 %	0.77 %
Q4	17.83 %	− 6.33 %	−1.36 %	− 0.73 %	− 1.04 %
Q5 (highest)	19.07 %	− 10.55 %	−1.91 %	− 1.76 %	− 1.83 %

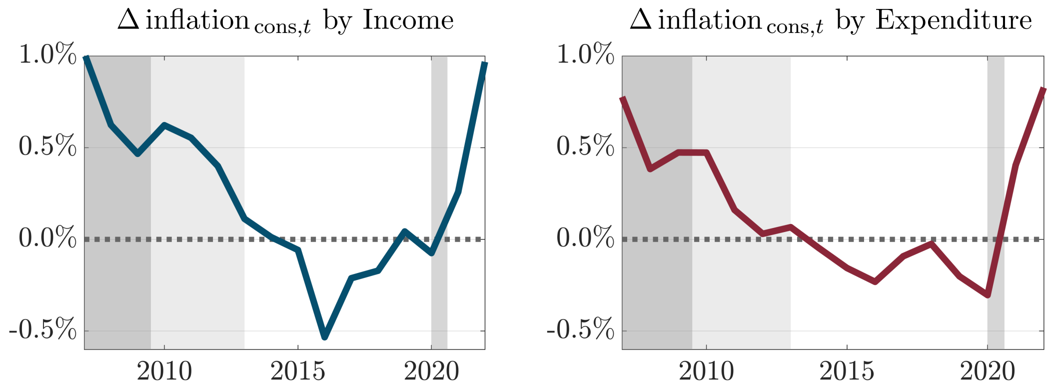
Welfare cost are the money-metric consumption-equivalent change $cv^{\text{total}} - 1$ in percent on spending. The Price share in these welfare cost uses the superlative Fisher price CEV. "Price-only" columns report the Laspeyres-, Paasche-, and Fisher-based spending boosts required to offset price movements alone.

5.6 Evidence on the mechanism

I show that relative price movements and spending reallocation in the Great Recession line up with the model's predictions. Treating quantitative gaps as recessionary movements in marginal costs allows me to gauge the markup channel's importance.

Differences in inflation rates. Abstracting from marginal-cost shifts, the model implies that markups faced by the poor rise relative to those faced by the rich in downturns. In the data (where prices also move with costs) relative prices borne by the poor nonetheless increase in recessions.

Figure 6: Inflation for Poor Households Is Higher in Recessions



Törnqvist inflation for poor minus rich households. Left: income-based segmentation. Right: expenditure-based segmentation. Indices reflect product choice only (search is held fixed by construction).

I compute Törnqvist indices separately for $h \in \{\text{poor}, \text{rich}\}$,

$$\text{inflation}_{h,t} \equiv \exp\left(\sum_{i \in \mathscr{I}} \frac{\text{share}_{i,h,t} + \text{share}_{i,h,t-1}}{2} \log\left(\frac{\text{price}_{i,t}}{\text{price}_{i,t-1}}\right)\right), \tag{26}$$

where $\text{share}_{i,h,t}$ is the average expenditure share of h on barcode i. Crucially, barcode-level inflation rates are averaged across households and therefore identical across segments; differences in $\text{inflation}_{h,t}$ arise only from composition, not search. I report results under two segmentations: (i) rich if income exceeds the region–time median; (ii) rich if NielsenIQ barcode spending is above the median.[5] Define the inflation gap

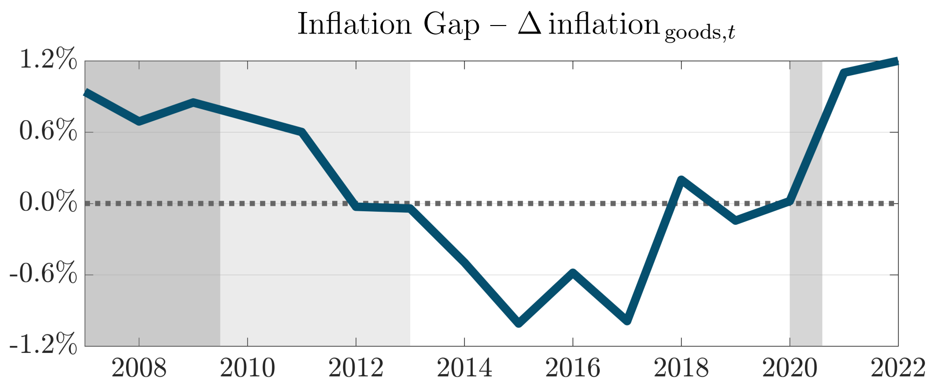
$$\Delta\,\text{inflation}_{\text{consumers},t} \equiv \text{inflation}_{\text{poor},t} - \text{inflation}_{\text{rich},t}. \tag{27}$$

Figure 6 shows that prices faced by the poor rise relative to those faced by the rich during and after recessions.[6]

[5]Income mitigates concerns about substitutes outside NielsenIQ (e.g., restaurants vs. groceries). Expenditures are informative given the absence of wealth data and possible mismeasurement from income alone.

[6]Recessions thus drive inflation inequality (cf. Jaravel, 2019); the long-run gap reflects only a partial reversal in expansions.

Figure 7: Inflation for Cheap Goods Is Higher in Recessions



Difference in Törnqvist inflation between cheap and premium goods, where the partition is based on time-averaged barcode-level premium scores.

In the model, cheaper (lower-quality) options become relatively more expensive in recessions. I verify this by computing Törnqvist indices for a partition $\{\mathscr{I}_k\}$ with $k \in \{\text{cheap}, \text{premium}\}$:

$$\text{inflation}_{k,t} \equiv \exp\left(\sum_{i \in \mathscr{I}_k} \frac{\text{share}_{i,t} + \text{share}_{i,t-1}}{2} \log\left(\frac{\text{price}_{i,t}}{\text{price}_{i,t-1}}\right)\right). \tag{28}$$

"Cheap" goods have sales-weighted region–time averages of premium_{irt} below zero; "premium" goods have averages above zero. Averaging premium scores over time prevents mechanical comovement with inflation.[7] I define

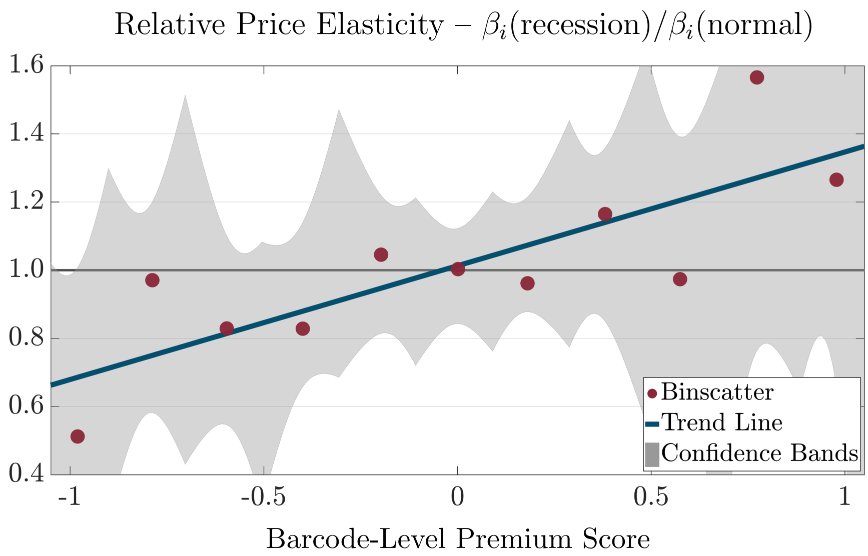
$$\Delta\,\text{inflation}_{\text{goods},t} \equiv \text{inflation}_{\text{cheap},t} - \text{inflation}_{\text{premium},t}. \tag{29}$$

Figure 7 shows that the relative price of cheap varieties rises in recessions and their aftermath, consistent with "cheapflation" documented internationally after COVID-19 (Cavallo and Kryvtsov, 2024).

Price elasticities during the Great Recession. The model predicts that affluent households become less price-sensitive to cheap goods when they real-

[7]This rules out spurious correlations from mean reversion in marginal costs.

Figure 8: The Rich Become Less Elastic Toward Cheap Goods in Recessions



Binscatter of premium scores vs. the ratio of recession to normal elasticities among higher-income households.

locate toward them in recessions. I estimate barcode-level elasticities for high-income households in and out of the Great Recession and compute the ratio $\beta_i(\text{recession})/\beta_i(\text{normal})$. Ratios below one indicate reduced sensitivity in the recession for barcode i. Figure 8 plots these ratios against barcode premium scores. Consistent with my model, affluent households became less elastic for cheap goods and more elastic for expensive ones.

6 Conclusion

Poor and rich households purchase different mixes of varieties within sectors, with poorer households allocating more to budget tiers. I show that recessionary demand reallocation toward inexpensive goods interacts with oligopolistic pricing in a way that systematically tilts relative prices. The paper brings two ingredients to this question: micro evidence on expenditure shares along the premium margin and the cross-sectional gradient of price sensitivity, and a calibrated model that matches these moments. Holding marginal costs and market structure fixed, feeding the observed distributional shifts from the Great Recession and COVID-19 isolates a demand-composition mechanism: shifting mass toward budget tiers weakens cross-tier competitive discipline from premium varieties, raising low-tier markups and moving relative prices against baskets purchased by the poor.

Quantitatively, the model yields large, tier-specific markup movements with muted aggregates. In the Great Recession, low-tier markups rise by 5.68 pp versus 2.30 pp at the high tier, increasing the low/high relative price by 2.59 %. In COVID-19, aggregate markups are essentially flat but split strongly (low +4.27 pp, high -2.71 pp), raising the low/high relative price by 5.27 %. These price tilts have first-order welfare consequences: for the bottom quintile, money-metric welfare losses exceed PCE-deflated spending by 4.92 pp in the Great Recession and 3.96 pp in COVID; prices account for 42.75 % of the bottom's loss in the Great Recession and mitigate about 10.5 % at the top in COVID.

Two implications follow. For measurement, a common price index obscures distributional inflation and basket-specific indices are necessary when demand is nonhomothetic and competition is tiered. For policy, cross-tier forces matter. Interventions that alter competitive pressure at premium margins (or shield necessity baskets) can materially change incidence even when aggregate markups move little.

The analysis is deliberately conservative: I study a static, partial-equilibrium model that holds costs, entry/exit, and quality investment fixed to isolate the demand-composition channel. Extending the framework to general equilibrium with income dynamics, endogenous product quality, and explicit supply disturbances would allow evaluation of counterfactual policies (targeted transfers, VAT changes, necessity subsidies) and refine distributional inflation measurement. The core lesson, however, is robust: when recessions push expenditure toward budget tiers, cross-tier discipline weakens, low-tier pricing power rises, and the resulting "cheapflation" loads disproportionately on poorer households.

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Table 6: Relative Price Changes and Segment Competition

	$\Delta \log(p_s^L/p_s^H)$			
	(1)	(2)	(3)	(4)
Great Recession				
Intercept	0.0249^{***}	0.0249^{***}	0.0249^{***}	0.0249^{***}
z-score HHI$_s$	0.0020			
z-score K-HHI$_s$		-0.0089^{***}		
z-score K-HHI$_s(p_s^L)$			0.0105^{***}	-0.0002
z-score K-HHI$_s(p_s^H)$			-0.0099^{***}	0.0009
z-score $\Delta\, x_s^{\text{bottom}}$				0.0273^{***}
z-score $\Delta\, x_s^{\text{top}}$				-0.0163^{***}
z-score Initial x_s^{bottom}				0.0073^{***}
z-score Initial x_s^{top}				-0.0109^{***}
R^2	0.025	0.461	0.539	0.997
COVID-19 Pandemic				
Intercept	0.0522^{***}	0.0522^{***}	0.0522^{***}	0.0522^{***}
z-score HHI$_s$	0.0064^{**}			
z-score K-HHI$_s$		-0.0021		
z-score K-HHI$_s(p_s^L)$			0.0049	-0.0048
z-score K-HHI$_s(p_s^H)$			0.0019	0.0014
z-score $\Delta\, x_s^{\text{bottom}}$				0.0246^{***}
z-score $\Delta\, x_s^{\text{top}}$				-0.0152^{**}
z-score Initial x_s^{bottom}				0.0147^{*}
z-score Initial x_s^{top}				-0.0103^{*}
R^2	0.277	0.029	0.254	0.986

Within-model sector regressions. Dependent variable is the percentage change in the low–high relative price. K-HHIs are computed from pre-recession shares and prices using a Gaussian kernel with Silverman h. Columns add regressors sequentially. All covariates are z-scored. Bottom/top cross-sector demand shifts are based in below and above median households in the expenditure distribution. Since regressors are z-scored, percentage point partial effects for regressor j in the main text are computed as $100 \times \big(\exp(\hat{\alpha} + \hat{\beta}_j \times 1) - \exp(\hat{\alpha})\big)$. Significance: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table 7: Variety-level markup changes, quality, and localized competition

	$\Delta \log \mu_{is}$				
	(1)	(2)	(3)	(4)	(5)
Great Recession					
Intercept	0.0331***	0.0207***	0.0207***	0.0207***	0.0207***
z-score HHI_s	0.0069*				
Low-Quality Dummy $\mathbb{1}\{i \in L\}$		0.0249***	0.0249***	0.0249***	0.0249***
z-score $\text{K-HHI}_s(p_s^L)$			0.00938**	0.00085	
z-score $\text{K-HHI}_s(p_s^H)$			−0.00419	0.00058	
z-score $\Delta\, x_s^{\text{bottom}}$				0.00842	−0.00825**
z-score $\Delta\, x_s^{\text{top}}$				0.01225	0.02321***
z-score $\Delta\, x_s^{\text{bottom}} \times \mathbb{1}\{i \in L\}$					0.03524***
z-score $\Delta\, x_s^{\text{top}} \times \mathbb{1}\{i \in L\}$					−0.02398***
z-score Initial x_s^{bottom}				−0.00328	−0.00231*
z-score Initial x_s^{top}				0.00289	0.00317***
R^2	0.077	0.262	0.361	0.930	0.992
COVID-19 Pandemic					
Intercept	0.0049	0.0212***	0.0212***	0.0212***	0.0212***
z-score HHI_s	0.0010				
Low-Quality Dummy $\mathbb{1}\{i \in L\}$		0.05219***	0.05219***	0.05219***	0.05219***
z-score $\text{K-HHI}_s(p_s^L)$			0.00239	0.00321	
z-score $\text{K-HHI}_s(p_s^H)$			−0.00223	0.00043	
z-score $\Delta\, x_s^{\text{bottom}}$				0.00769	0.00031
z-score $\Delta\, x_s^{\text{top}}$				0.01090	0.01258***
z-score $\Delta\, x_s^{\text{bottom}} \times \mathbb{1}\{i \in L\}$					0.02130***
z-score $\Delta\, x_s^{\text{top}} \times \mathbb{1}\{i \in L\}$					−0.01069***
z-score Initial x_s^{bottom}				−0.00955	−0.00388***
z-score Initial x_s^{top}				0.00569	0.00263*
R^2	0.001	0.675	0.679	0.965	0.997

Within-model variety regressions. Column (1) uses HHI only. Column (2) includes only a quality fixed effect. Column (3) adds localized competition. Column (4) further adds cross-sector demand shifts at the bottom and top (below and above median) of the expenditure distribution. Column (5) replaces localized competition with the interactions of expenditure-specific cross-sector spending shifts with the low-quality indicator. All listed covariates are standardized (z-scores). Since regressors are z-scored, percentage point partial effects for regressor j are computed as $100 \times \big(\exp(\hat{\alpha} + \hat{\beta}_j \times 1) - \exp(\hat{\alpha})\big)$ with baseline $\hat{\alpha}$. Significance: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Appendix A - Additional Results

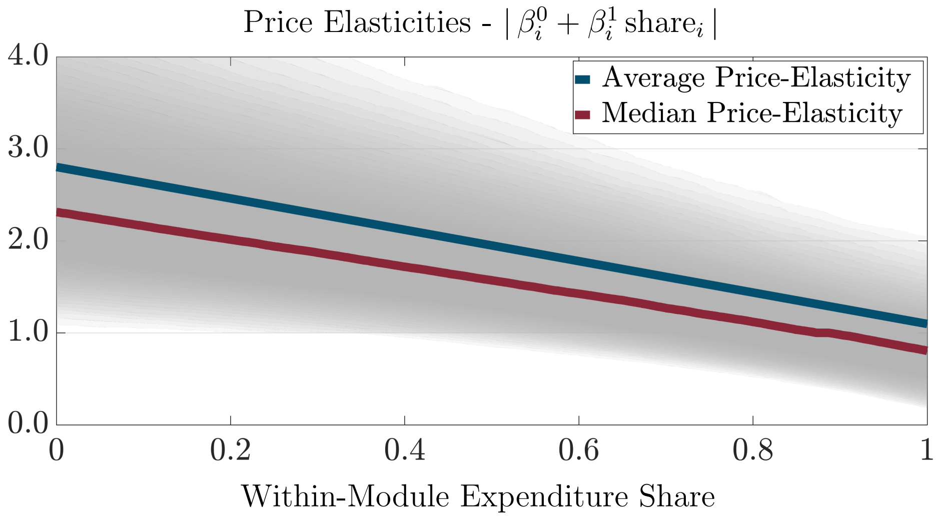
Fact 3 via stratification of population. To establish fact 3 from Section 3 in the data, here, I estimate barcode-specific price elasticities as a function of within-module expenditure shares. Specifically, I conduct barcode-level instrumental variable regressions to estimate a log-linearization of my demand system from (16). My approach leverages the fact that model-implied price elasticities depend on household characteristics only through expenditure shares. In particular, I interact household h's expenditure share for variety i within product module $m(i)$ with household-specific log prices. The resulting regression equation is:

$$\begin{aligned}\log \text{quantity}_{iht} = \; & \alpha_{ih} + \alpha_{it} + \alpha_{ir(h)} + \left(\beta_i^0 + \beta_i^1 \,\text{share}_{iht}\right) \log \text{price}_{iht} \qquad (30) \\ & + \sum\nolimits_{j \in \mathscr{C}_{iht}} \beta_{ij} \log \text{price}_{jr(h)t} + \gamma_i \log \text{expenditure}_{ht} + \epsilon_{iht}.\end{aligned}$$

In this regression, $\beta_i(x) \equiv \left| \beta_i^0 + \beta_i^1 \, x \right|$ can be interpreted as barcode i's own price elasticity among consumers who allocate an expenditure shares x on barcode i in product module $m(i)$. The regression controls for both household expenditures and a carefully constructed set of household-specific competitors for each barcode i. Using data on shopping trips from the consumer panel and store-level pricing information from NielsenIQ's Retail Scanner data, I ensure that $\mathscr{C}_{iht}$ is comprised of barcodes $j \in m(i)$ that are actually available to household h at $\text{price}_{jr(h)t}$.

In order to address potential endogeneity issues in the relationship between quantity_{iht} and price_{iht}, note that idiosyncratic determinants of the quantity choice of a particular household are likely orthogonal to retail prices. Therefore, the reverse causality concern boils down to the presence of local demand shocks that are observable to retailers (and thus reflected in pricing) but unobservable to the econometrician. To address this concern, I construct a set price instruments in the spirit of Hausman (1996). Specifically, I instrument price_{iht} with the economy-wide average price for barcode i in year t, excluding observations from region $r(h)$. Moreover, since consumption choices along the premium margin are highly correlated with income, I instrument expenditure shares with household income_{ht}. Price elasticities are, therefore, identified by within-household/region/time price variation that is explained by economy-wide price movements. The identifying assumption is that pricing decisions that apply to barcodes throughout the nation are orthogonal to local demand conditions. This exclusion restriction is broadly consistent

Figure 9: Cross-sectional Distribution of Price Elasticities



This graph depicts the distribution of $i \mapsto \beta_i(x)$ across the universe of NielsenIQ barcodes. The graph shows that both average and median price elasticities decrease as within-module spending shares x increase. The shaded area extends from the 5th to 95th percentile of the cross-sectional distribution.

with evidence of uniform pricing documented by DellaVigna and Gentzkow (2019). Since all coefficients in (30) are indexed at the barcode level, under mild clustering conditions on ϵ_{iht}, all regressions are run barcode by barcode.

Figure 9 illustrates the distribution of $i \mapsto \beta_i(x)$ across the universe of NielsenIQ barcodes. The graph shows that both average and median price elasticities decrease as within-module spending shares x increase. That is, as consumers increase their relative spending on a particular variety, they become less price-elastic with respect to it. Consequently, from Figure 9, we can deduce that poor consumers, who oftentimes concentrate their within-sector spending on the least expensive options available, inelastically choose these low-cost alternatives. Similarly, wealthier consumers are inelastic regarding premium goods, as they devote a significant portion of their spending to these pricier varieties. By contrast, middle-class consumers, whose spending is more evenly distributed across a range of varieties along the premium margin, display greater price elasticity in either direction.

When linking this empirical observation to model-implied price elasticities, note that in a homothetic model environment, expenditure shares are uniform across the entire population. Consequently, there is no cross-sectional heterogeneity in price elasticities. However, as shown in Figure 1, nonhomothetic preferences lead to variation in expenditure shares on different-quality products across the expenditure distribution. Figure 2 further illustrates that model-implied price elasticities generally decrease as expenditure shares increase.

A remark is in order: While the price elasticities in Figure 9 align with other estimates based on NielsenIQ data, such as those from Hitsch, Hortacsu, and Lin (2019), they are significantly lower than standard macro estimates of elasticities of substitution. One possible explanation for this discrepancy is that elasticities of substitution within grocery stores are, in fact, relatively low, and that retail markups are largely influenced by competition among stores. The key insight from this exercise is qualitative: when facing a fixed product assortment, households become less price elastic for a particular variety as their expenditure share on it increases.

Appendix B - Model and Derivations

Households' program. Consumers take as given the economy's price vector $\boldsymbol{p} = \left(\boldsymbol{p}_s : s \in \mathscr{S}\right)$ where $\boldsymbol{p}_s = \left(p_{iqs} : q \in \mathscr{Q} \text{ and } i \in \mathscr{I}_{qs}\right) \ \forall \ s \in \mathscr{S}$. They choose allocations $\left\{c_{iqs}\right\}$ to minimize the expenditure necessary to attain real consumption c. That is, they solve

$$\inf_{\{c_{iqs},\, c_s\}} \int_{\mathscr{S}} \sum_{q=1}^{Q} \sum_{i=1}^{N_{qs}} p_{iqs} c_{iqs} \, ds \tag{31}$$

subject to

$$\int_{\mathscr{S}} \left(\frac{c_s}{c}\right)^{\frac{\eta-1}{\eta}} ds = 1 \tag{32}$$

and

$$\sum_{q=1}^{Q} \sum_{i=1}^{N_{qs}} \left(\frac{\varphi_q}{c_s^{(1-\sigma)(1-\xi_q)}}\right)^{\frac{1}{\sigma}} \left(\frac{c_{iqs}}{c_s}\right)^{\frac{\sigma-1}{\sigma}} = 1 \quad \forall \quad s \in \mathscr{S}. \tag{33}$$

Intermediate Hicksian demand. Invoking a standard separation theorem, the first step is to minimize within-sector expenditures to attain a given level of sectoral consumption c_s. To that end, households solve

$$\min_{\{c_{iqs}\}} \left\{ \sum_{q=1}^{Q} \sum_{i=1}^{N_{qs}} p_{iqs} c_{iqs} \;\middle|\; \sum_{q=1}^{Q} \sum_{i=1}^{N_{qs}} \varphi_q^{\frac{1}{\sigma}} \left(\frac{c_{iqs}}{c_s^{\xi_q}}\right)^{\frac{\sigma-1}{\sigma}} = 1 \right\}. \tag{34}$$

With λ being the Lagrange-multiplier on the nonhomothetic CES aggregator, the first-order conditions with respect to c_{iqs} are

$$p_{iqs} = \lambda \, \frac{\sigma-1}{\sigma} \, \varphi_q^{\frac{1}{\sigma}} \left(\frac{c_{iqs}}{c_s^{\xi_q}}\right)^{\frac{\sigma-1}{\sigma}-1} \frac{1}{c_s^{\xi_q}}. \tag{35}$$

Multiplying by c_{iqs} and summing over producers and quality-bins we obtain sectoral spending y_s as

$$y_s \equiv \sum_{q=1}^{Q} \sum_{i=1}^{N_{qs}} p_{iqs} c_{iqs} = \lambda \, \frac{\sigma-1}{\sigma} \tag{36}$$

where the latter equality follows from the definition of the nonhomothetic CES aggregator in (33). Sectoral expenditure shares are, thus, given as

$$x_{iqs} \equiv \frac{p_{iqs} c_{iqs}(c_s, \boldsymbol{p}_s)}{\sum_{q=1}^{Q} \sum_{i=1}^{N_{qs}} p_{iqs} c_{iqs}} = \varphi_q^{\frac{1}{\sigma}} \left(\frac{c_{iqs}}{c_s^{\xi_q}} \right)^{\frac{\sigma-1}{\sigma}}. \tag{37}$$

At this point, we can define nonhomothetic ideal price-index $p_s(c_s, \boldsymbol{p}_s)$ to satisfy

$$y_s = p_s(c_s, \boldsymbol{p}_s)\, c_s = \sum_{q=1}^{Q} \sum_{i=1}^{N_{qs}} p_{iqs} c_{iqs}. \tag{38}$$

It is instructive to use (36) and (38) and rearrange terms in (35) to obtain

$$c_{iqs} = \varphi_q \left(\frac{p_{iqs}}{p_s(c_s, \boldsymbol{p}_s)} \right)^{-\sigma} c_s^{\sigma + (1-\sigma)\, \xi_q}. \tag{39}$$

Defining the nonhomothetic taste-shifter $\psi_q(c_s)$ as

$$\psi_q(c_s) \equiv \varphi_q\, c_s^{(1-\sigma)(\xi_q - 1)} \tag{40}$$

intermediate Hicksian demand functions are, in turn, concisely written as

$$c_{iqs}(c_s, \boldsymbol{p}_s) = \psi_q(c_s) \left(\frac{p_{iqs}}{p_s(c_s, \boldsymbol{p}_s)} \right)^{-\sigma} c_s. \tag{41}$$

To obtain an expression for $p_s(c_s, \boldsymbol{p}_s)$, expenditure shares can be rewritten as

$$x_{iqs} = \varphi_q \left(\frac{p_{iqs}}{p_s(c_s, \boldsymbol{p}_s)} \right)^{1-\sigma} c_s^{(1-\sigma)(\xi_q - 1)} \tag{42}$$

Since $\sum_{q=1}^{Q} \sum_{i=1}^{N_{qs}} x_{iqs} = 1$, it then follows that

$$p_s(c_s, \boldsymbol{p}_s) = \left(\sum_{q=1}^{Q} \sum_{i=1}^{N_{qs}} \varphi_q\, p_{iqs}^{1-\sigma} c_s^{(1-\sigma)(\xi_q - 1)} \right)^{\frac{1}{1-\sigma}}. \tag{43}$$

Intermediate Marshallian demand. Marshallian demand is not available in closed-form. We can, however, recover the sectoral nonhomothetic ideal price-index $p_s(y_s, \boldsymbol{p}_s)$ as a function of sectoral spending y_s in a single fixed-point problem. That is, in a slight abuse of notation,

$$p_s(y_s, \boldsymbol{p}_s) = \text{fix}\left\{ p \mapsto \left(\sum_{q=1}^{Q} \sum_{i=1}^{N_{qs}} \varphi_q \, p_{iqs}^{1-\sigma} \left(\frac{y_s}{p} \right)^{(1-\sigma)(\xi_q - 1)} \right)^{\frac{1}{1-\sigma}} \right\}. \tag{44}$$

This object, in turn, pins down intermediate Marshallian demand as

$$c_{iqs}(y_s, \boldsymbol{p}_s) = \varphi_q \left(\frac{p_{iqs}}{p_s(y_s, \boldsymbol{p}_s)} \right)^{-\sigma} \left(\frac{y_s}{p_s(y_s, \boldsymbol{p}_s)} \right)^{\sigma + (1-\sigma)\,\xi_q}. \tag{45}$$

Demand for sectoral aggregates. Conditional on the optimal within-sector allocation of consumption, we can next determine the expenditure-minimizing allocation of sectoral real consumption indices. Note that the nonhomothetic ideal price-index recovered in (43) encapsulates optimality of the consumers' within-sector decision problem. As a consequence, at the outer nest, we can think of the households' program as a homothetic expenditure minimization problem with a non-linear pricing structure. That is,

$$\inf_{\{c_s\}} \left\{ \int_{\mathscr{S}} p_s(c_s, \boldsymbol{p}_s)\, c_s \, ds \,\middle|\, \int_{\mathscr{S}} \left(\frac{c_s}{c} \right)^{\frac{\eta-1}{\eta}} ds = 1 \right\}. \tag{46}$$

With λ being the Lagrange-multiplier on the across-sector CES aggregator, the first-order conditions with respect to c_s are such that

$$\frac{\partial\, p_s(c_s, \boldsymbol{p}_s)}{\partial\, c_s}\, c_s + p_s(c_s, \boldsymbol{p}_s) = \lambda\, \frac{\eta - 1}{\eta} \left(\frac{c_s}{c} \right)^{\frac{\eta-1}{\eta} - 1} \frac{1}{c} \tag{47}$$

To arrive at the first-order condition from (12), first off, we recover the partial derivative of sectoral prices with respect to real consumption as

$$\frac{\partial\, p_s(c_s, \boldsymbol{p}_s)}{\partial\, c_s} = \frac{\partial}{\partial\, c_s}\left[\left(\sum_{q=1}^{Q}\sum_{i=1}^{N_{qs}} \varphi_q\, p_{iqs}^{1-\sigma}\, c_s^{(1-\sigma)(\xi_q-1)}\right)^{\frac{1}{1-\sigma}}\right] \tag{48}$$

$$= p_s(c_s, \boldsymbol{p}_s)^{\sigma} \sum_{q=1}^{Q}\sum_{i=1}^{N_{qs}} \varphi_q\, p_{iqs}^{1-\sigma}\, c_s^{(1-\sigma)(\xi_q-1)-1}(\xi_q - 1) \tag{49}$$

Using (42) and rewriting the derivative as an elasticity, the above expression simplifies to

$$\frac{\partial\, \log p_s(c_s, \boldsymbol{p}_s)}{\partial\, \log c_s} = \sum_{q=1}^{Q}\sum_{i=1}^{N_{qs}} x_{iqs}(c_s, \boldsymbol{p}_s)\, \xi_q - 1. \tag{50}$$

For notational compactness, I define

$$\bar{\xi}_s(c_s, \boldsymbol{p}_s) = \sum_{q=1}^{Q}\sum_{i=1}^{N_{qs}} x_{iqs}(c_s, \boldsymbol{p}_s)\, \xi_q. \tag{51}$$

Multiplying the first-order condition from (47) by c_s we obtain

$$\left[\frac{\partial\, \log p_s(c_s, \boldsymbol{p}_s)}{\partial\, \log c_s} + 1\right] p_s(c_s, \boldsymbol{p}_s)\, c_s = \lambda\, \frac{\eta - 1}{\eta}\left(\frac{c_s}{c}\right)^{\frac{\eta-1}{\eta}}. \tag{52}$$

It, thus, follows that

$$p_s(c_s, \boldsymbol{p}_s)\, c_s\; \bar{\xi}_s(c_s, \boldsymbol{p}_s) = \lambda\, \frac{\eta - 1}{\eta}\left(\frac{c_s}{c}\right)^{\frac{\eta-1}{\eta}}. \tag{53}$$

We can now integrate over sectors and use the definition of across-sector aggregation from (32) to see that

$$\int_{\mathscr{S}} p_s(c_s, \boldsymbol{p}_s)\, c_s\; \bar{\xi}_s(c_s, \boldsymbol{p}_s)\, ds = \lambda\, \frac{\eta - 1}{\eta}. \tag{54}$$

It then immediately follows that

$$\frac{p_s(c_s, \boldsymbol{p}_s)\, c_s\; \bar{\xi}_s(c_s, \boldsymbol{p}_s)}{\int p_s(c_s, \boldsymbol{p}_s)\; c_s\; \bar{\xi}_s(c_s, \boldsymbol{p}_s)\, ds} = \left(\frac{c_s}{c}\right)^{\frac{\eta-1}{\eta}}. \tag{55}$$

Using the fact that $p_s(c_s, \boldsymbol{p}_s)\, c_s = \sum_q \sum_i p_{iqs}\, c_{iqs}(c_s, \boldsymbol{p}_s)$, this can be rewritten as

$$\frac{\sum_q \sum_i p_{iqs}\, c_{iqs}(c_s, \boldsymbol{p}_s)\, \xi_q}{\int \sum_q \sum_i p_{iqs}\, c_{iqs}(c_s, \boldsymbol{p}_s)\, \xi_q\, ds} = \left(\frac{c_s}{c}\right)^{\frac{\eta-1}{\eta}}. \tag{56}$$

Bertrand competition. To obtain an expression for the Bertrand price-elasticity of variety (k, r, s), I, first off, take the logarithm of the intermediate Hicksian demand function from (41). That is,

$$\log c_{krs}(c_s, \boldsymbol{p}_s) = \log \varphi_r - \sigma \log p_{krs} + \sigma \log p_s(c_s, \boldsymbol{p}_s) + \big(\sigma + (1-\sigma)\, \xi_r\big)\, \log c_s. \tag{57}$$

The partial derivative with respect to log p_{krs} is then given as

$$\frac{\partial \log c_{krs}(c_s, \boldsymbol{p}_s)}{\partial \log p_{krs}} = -\sigma + \sigma \frac{\partial \log p_s(c_s, \boldsymbol{p}_s)}{\partial \log p_{krs}} + \Big(\sigma + (1-\sigma)\, \xi_q\Big) \frac{\partial \log c_s}{\partial \log p_{krs}}. \tag{58}$$

In order to see how the nonhomothetic ideal price-index p_s responds to a *ceteris paribus* change in p_{krs}, I compute

$$\begin{aligned} \frac{\partial p_s(c_s, \boldsymbol{p}_s)}{\partial p_{krs}} &= \frac{\partial}{\partial p_{krs}} \left[\left(\sum_{q=1}^{Q} \sum_{i=1}^{N_{qs}} \varphi_q\, p_{iqs}^{1-\sigma}\, c_s^{(1-\sigma)(\xi_q - 1)} \right)^{\frac{1}{1-\sigma}} \right] \\ &= p_s(c_s, \boldsymbol{p}_s)^{\sigma}\, \varphi_q\, p_{iqs}^{-\sigma}\, c_s^{(1-\sigma)(\xi_q-1)-1} \\ &\quad + p_s(c_s, \boldsymbol{p}_s)^{\sigma} \frac{\partial c_s}{\partial p_{krs}} \sum_{q=1}^{Q} \sum_{i=1}^{N_{qs}} \varphi_q\, p_{iqs}^{1-\sigma}\, c_s^{(1-\sigma)(\xi_q-1)-1} (\xi_q - 1) \end{aligned} \tag{59}$$

Using (42) and writing the above derivative as an elasticity, we obtain

$$\frac{\partial \log p_s(c_s, \boldsymbol{p}_s)}{\partial \log p_{kqs}} = x_{kqs}(c_s, \boldsymbol{p}_s) + \frac{\partial \log c_s}{\partial \log p_{kqs}} \Big[\bar{\xi}_s(c_s, \boldsymbol{p}_s) - 1 \Big] \tag{60}$$

Next, $\partial \log c_s / \partial \log p_{kqs}$ is most conveniently recovered in an application of the implicit function theorem. Specifically, I define

$$F(c_s, p_{kqs}) \equiv \sum_{q=1}^{Q} \sum_{i=1}^{N_{qs}} p_{iqs} c_{iqs}(c_s, \boldsymbol{p}_s)\, \xi_q - A \left(\frac{c_s}{c}\right)^{\frac{\eta-1}{\eta}} \tag{61}$$

We can, therefore compute

$$\begin{aligned}
\frac{\partial F(c_s, p_{kqs})}{\partial c_s} = {} & \sigma \sum_{q=1}^{Q} \varphi_q\, \xi_q \sum_{i=1}^{N_{qs}} \left(\frac{p_{iqs}}{p_s(c_s, \boldsymbol{p}_s)}\right)^{1-\sigma} \frac{\partial p_s(c_s, \boldsymbol{p}_s)}{\partial c_s}\, c_s^{\sigma+(1-\sigma)\,\xi_q} \\
& + \sum_{q=1}^{Q} \varphi_q\, \xi_q \left(\sigma + (1-\sigma)\, \xi_q\right) \sum_{i=1}^{N_{qs}} \left(\frac{p_{iqs}}{p_s(c_s, \boldsymbol{p}_s)}\right)^{1-\sigma} p_s(c_s, \boldsymbol{p}_s)\, c_s^{(1-\sigma)(\xi_q-1)} \\
& - A\, \frac{\eta-1}{\eta} \left(\frac{c_s}{c}\right)^{\frac{\eta-1}{\eta}-1}
\end{aligned} \tag{62}$$

as well as

$$\begin{aligned}
\frac{\partial F(c_s, p_{kqs})}{\partial p_{krs}} = {} & \varphi_r\, \xi_r\, (1-\sigma) \left(\frac{p_{krs}}{p_s(c_s, \boldsymbol{p}_s)}\right)^{-\sigma} c_s^{\sigma+(1-\sigma)\,\xi_r} \\
& + \sigma \sum_{q=1}^{Q} \varphi_q\, \xi_q \sum_{i=1}^{N_{qs}} \left(\frac{p_{iqs}}{p_s(c_s, \boldsymbol{p}_s)}\right)^{1-\sigma} \left.\frac{\partial p_s(c_s, \boldsymbol{p}_s)}{\partial p_{krs}}\right|_{c_s} c_s^{\sigma+(1-\sigma)\,\xi_q}
\end{aligned} \tag{63}$$

Note that, in the partial derivative of F with respect to p_{krs} sectoral consumption c_s is not responsive to price changes. That is,

$$\left.\frac{\partial p_s(c_s, \boldsymbol{p}_s)}{\partial p_{krs}}\right|_{c_s} \neq \frac{\partial p_s(c_s, \boldsymbol{p}_s)}{\partial p_{krs}} \tag{64}$$

and, specifically,

$$\left.\frac{\partial p_s(c_s, \boldsymbol{p}_s)}{\partial p_{krs}}\right|_{c_s} = x_{krs}(c_s, \boldsymbol{p}_s). \tag{65}$$

By the implicit function theorem (and a few straightforward algebraic manipulations), we then have

$$\frac{\partial \log c_s}{\partial \log p_{krs}} = -\frac{\dfrac{\partial F(c_s, p_{krs})}{\partial p_{krs}} \dfrac{p_{krs}}{p_s(c_s, \boldsymbol{p}_s)}}{\dfrac{\partial F(c_s, p_{krs})}{\partial p_{krs}} \dfrac{c_s}{p_s(c_s, \boldsymbol{p}_s)}} \tag{66}$$

Substituting in the results from above and simplifying the expression we find that

$$\frac{\partial \log c_s}{\partial \log p_{krs}} = -\frac{(1-\sigma)\, x_{krs}(c_s, \boldsymbol{p}_s)\, \xi_r + \sigma\, x_{krs}(c_s, \boldsymbol{p}_s)\, \bar{\xi}_s(c_s, \boldsymbol{p}_s)}{(1-\sigma)\, \bar{\xi}^2_s(c_s, \boldsymbol{p}_s) + \sigma\, \bar{\xi}_s(c_s, \boldsymbol{p}_s)^2 - \dfrac{\eta - 1}{\eta}\, \bar{\xi}_s(c_s, \boldsymbol{p}_s)}. \tag{67}$$

Finally, substituting (60) and (67) into (58) and by duality, the price-elasticity of variety (k, r, s) is given as

$$\left| \frac{\partial \log c_{krs}(y, \boldsymbol{p})}{\partial \log p_{krs}} \right| = \big(1 - x_{krs}(y, \boldsymbol{p})\big)\, \sigma + x_{krs}(y, \boldsymbol{p})\, \eta\, \zeta_{krs}(y, \boldsymbol{p}) \tag{68}$$

where

$$\zeta_{krs}(y, \boldsymbol{p}) \equiv \frac{\Big(\sigma\, \bar{\xi}(y, \boldsymbol{p}) + (1-\sigma)\, \xi_r\Big)^2}{\sigma\, \eta\, \bar{\xi}(y, \boldsymbol{p})^2 + (1-\sigma)\, \eta\, \bar{\xi}^2(y, \boldsymbol{p}) + (1-\eta)\, \bar{\xi}(y, \boldsymbol{p})}. \tag{69}$$

Algorithm to compute Nash equilibrium.

1. For each $y \in \text{support}(g)$ guess $\{y_s^{(0)}\}$ such that $\int y_s^{(0)}\, ds = y$
2. In iteration n, conditional on $\{y_s^{(n)}\}$, find $\{\boldsymbol{\mu}_s^{(n)}\}$ such that equations (4) hold for

   $$\boldsymbol{p}_s^{(n)} = \boldsymbol{\mu}_s^{(n)} \circ \boldsymbol{\lambda}_s$$

   This is an isolated $Q \times N$ dimensional root-finding problem for each $s \in \mathscr{S}$.
3. Compute $c_{iqs}^{(n)} = c_{iqs}(y_s^{(n)}, \boldsymbol{\mu}_s^{(n)} \circ \boldsymbol{\lambda}_s)$.
4. Find $\{y_s^{(n+1)}\}$ and $c^{(n+1)}$ such that

$$\frac{\sum_q \sum_i \mu_{iqs}^{(n)}\, \lambda_{iqs}\, c_{iqs}^{(n)}\, \xi_q}{\int \sum_q \sum_i \mu_{iqs}^{(n)}\, \lambda_{iqs}\, c_{iqs}^{(n)}\, \xi_q\, ds} = \left(\frac{y_s^{(n+1)}}{p_s^{(n)}\; c^{(n+1)}} \right)^{\frac{\eta-1}{\eta}} \qquad \forall \quad s \in \mathscr{S}$$

and

$$\int y_s^{(n+1)}\, ds = y$$

5. Return to 2. and iterate until $\left\| \boldsymbol{\mu}_s^{(n+1)} - \boldsymbol{\mu}_s^{(n)} \right\| <$ tolerance.