Pass-Through and the Unequal Incidence of Commodity Shocks

Kunal Sangani

Harvard*

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Abstract

Empirical studies of commodity cost pass-through find that pass-through is incomplete: a 10 percent increase in costs causes downstream prices to rise less than 10 percent, even at long horizons. Using microdata from gasoline and food products, I find that incomplete pass-through in percentages often disguises *complete pass-through in levels*: a $1/unit increase in commodity costs leads to $1/unit higher downstream prices. Pass-through appears incomplete in percentages due to an additive margin between marginal costs and prices. A model in which firms bound the risk of profits falling short of overhead costs can account for this pricing behavior. Due to complete pass-through in levels, rising commodity costs lead to higher inflation rates for low-margin products in a category, though absolute price changes are similar across products. This generates cyclical inflation inequality. I document that food-at-home inflation is 10 percent more sensitive to upstream costs for households in the lowest income quintile than the highest. From 2020–2023, unequal commodity cost pass-through is responsible for two-thirds of the gap in food-at-home inflation rates experienced by low- and high-income households.

1   Introduction

Studies of how changes in commodity costs propagate downstream typically find evidence of *long-run incomplete pass-through*: when commodity prices increase 10 percent, downstream prices rise less than 10 percent (e.g., Kim and Cotterill 2008; Leibtag 2009; Nakamura and Zerom 2010; Hong and Li 2017). Pass-through remains incomplete even at long horizons and after accounting for the cost share of commodity inputs.

The leading explanation attributes incomplete pass-through to the shape of demand curves. When demand curves have a positive super-elasticity—i.e., the elasticity of demand increases with price—firms partly absorb increases in cost by reducing their markups. Theoretical work suggests myriad factors that can affect the super-elasticity of demand, such as market power (e.g., Atkeson and Burstein 2008), the extent of consumer heterogeneity (e.g., Mongey and Waugh 2023), or the underlying shape of consumer preferences (e.g., Klenow and Willis 2016). As a result, predicting the extent of commodity cost pass-through in practice requires estimating rich models of demand.

This paper studies the pass-through of commodity costs to downstream prices in a sample of markets where the amount of the commodity input required to produce downstream goods can be measured precisely. These markets, which include retail gasoline and several food products, allow me to measure pass-through both on a percentage basis—as pass-through is typically measured in the macro and international literatures—and on an absolute (dollars-and-cents) basis.

Nearly all of the studied markets exhibit *complete pass-through in levels*. That is, a $1/unit increase in commodity costs leads to a $1/unit increase in downstream prices. Complete pass-through in levels explains why pass-through measured in percentages appears incomplete: when price is greater than marginal cost, a $1/unit increase represents a smaller percentage increase in price than in marginal cost.

Microdata from retail gasoline and food products suggests that complete pass-through in levels explains not only the presence of incomplete "log pass-through," but also cross-sectional heterogeneity in log pass-through across firms or products in a market. For example, I find that gas stations with higher markups and food products within a category with higher retail markups have lower log pass-through, but no systematic difference in pass-through in levels. Intuitively, higher markups widen the gap between price and marginal cost, increasing the difference between the percentage changes in price and costs when absolute changes are identical.

Why do firms' prices exhibit complete pass-through in levels? In principle, the shape of demand curves could explain complete pass-through in levels if changes in the elasticity

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of demand along the demand curve lead firms to adjust their percentage markups in a way that coincides with complete pass-through in levels.[1] This is the case if the super-elasticity of demand curves is exactly equal to one.[2] At first blush, this explanation appears unlikely to justify complete pass-through in levels across markets, since estimated demand systems find variable super-elasticities of demand across products within a market, let alone across markets. Moreover, I find that super-elasticities of demand in my data, estimated using the technique developed by Burya and Mishra (2023), fall short of the magnitude required to explain pass-through in levels.

I thus propose a model where firms exhibit complete pass-through in levels despite facing isoelastic demand curves. In the model, firms have fixed overhead costs as well as variable costs for non-commodity inputs, which are bundled with the commodity into an output good sold to consumers. When firms set multiplicative markups over marginal cost, their variable profits scale with commodity costs, resulting in higher per-unit margins and entry when commodity costs are high and lower per-unit margins and exit when commodity costs are low. This variation in unit margins can be costly to firms. For example, in the presence of uncertainty over the level of demand, firms face a heightened risk of being unable to cover overhead expenses when commodity costs are low. When managers are averse to variable profits falling below overhead costs, firms can instead opt for an additive margin over marginal costs that is proportional to overhead costs. I show that a form of "safety margin constraint" pricing first discussed by Fellner (1948) (and subsequently explored by Day et al. 1971 and Altomonte et al. 2015) yields this pricing behavior. Since firms set an additive margin over marginal costs, changes in marginal costs are reflected one-for-one in changes in the output price.

This model generates additional predictions about the dynamics of entry and gross margins that I test in the data. While the standard multiplicative markup model predicts that entry rises with commodity costs while gross margins (variable profits as a share of sales) stay fixed, a model with safety margins predicts that entry is unrelated to commodity costs and gross margins fall when commodity costs rise. Using microdata from the markets described above, as well as broader data that covers several manufacturing sectors, I find

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[1]Another possible explanation is perfect competition: firms price at marginal cost and hence changes in marginal costs are reflected one-for-one in changes in prices. Perfect competition is the case where local costs (i.e., non-commodity input costs) account for the entirety of the gap between prices and commodity costs. Local costs are an important explanation for incomplete pass-through of exchange rate shocks (e.g., Burstein et al. 2003, Corsetti and Dedola 2005, Burstein et al. 2006). However, the limiting case of perfect competition is at odds with several features of the data: sluggish price adjustment, price dispersion for identical products, finite firm-level demand elasticities, and evidence of substantial markups over costs.

[2]The super-elasticity of demand is one when demand curves are semilog, i.e., $\log D(p) = -ap + b$. The fact that semilog demand curves generate complete pass-through in levels is discussed by Bulow and Pfleiderer (1983), Weyl and Fabinger (2013), and Mrázová and Neary (2017).

Figure 1: Within-category inflation inequality: Differences in coffee inflation for households in lowest vs. highest income quintiles track coffee commodity prices.
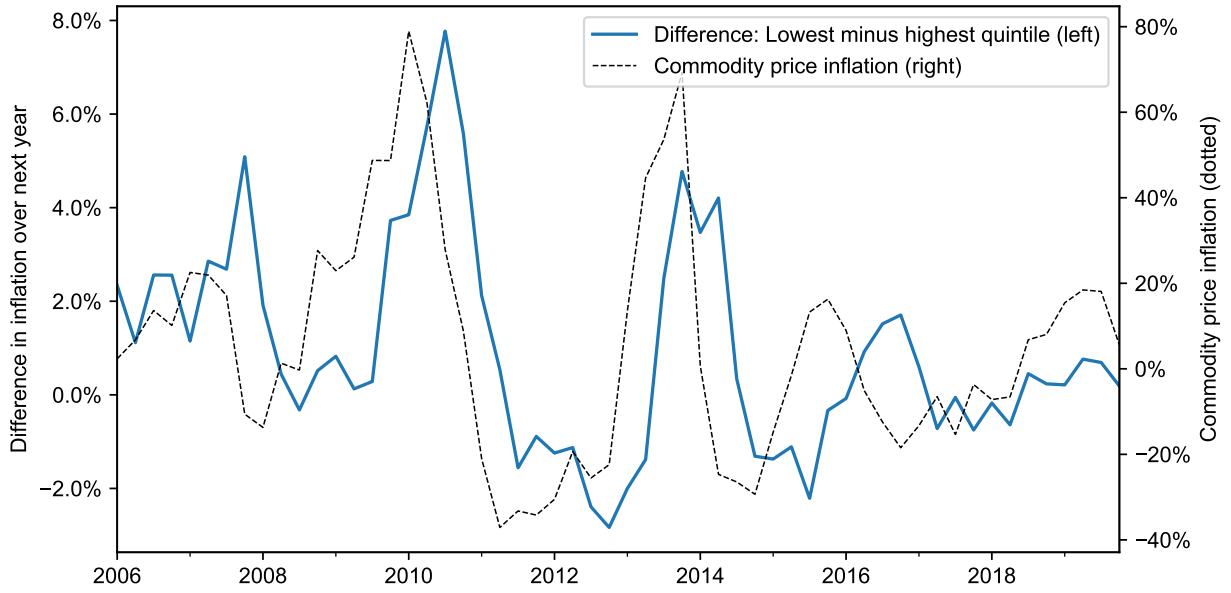
that the behavior of entry and gross margins in the data are consistent with the safety margins model rather than the standard model.

Complete pass-through in levels offers predictions for the extent of commodity cost pass-through that do not require estimating rich models of demand. In the final section of the paper, I demonstrate the relevance of these predictions by considering the implications of this pricing behavior for inflation inequality.

Specifically, I document a new, cyclical component of inflation inequality that arises due to complete pass-through in levels. Under complete pass-through in levels, low-margin products within a product category will exhibit higher inflation rates when commodity costs rise, even though absolute price changes are similar across products. Since low-income households tend to purchase lower-price and lower-markup products (Sangani 2022), rising commodity costs cause low-income households to face higher inflation rates even within narrow product categories. For example, as shown in Figure 1, the gap in coffee inflation rates experienced by low- and high-income households surges when coffee commodity prices are rising (e.g., in 2011 and 2014) and falls—even becoming negative—when commodity prices are falling (e.g., in 2012).

Aggregating over the food-at-home bundle, this pass-through behavior results in higher food-at-home inflation volatility for low-income households. By combining panelist-recorded purchases with store scanner data, I estimate income group-specific food-at-

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home price indices from 2006 to 2020. The log pass-through of upstream food price indices to food-at-home inflation for households in the lowest income quintile is 10 percent higher than for the highest income quintile, and the variance of annual food-at-home inflation rates experienced by the lowest income quintile is 20 percent greater than that of the highest income group.

I apply these estimates to changes in the food-at-home price level experienced by households since the onset of the COVID-19 pandemic in 2020. I estimate that food-at-home prices for the lowest income quintile grew 15.6pp from 2020–2023, compared to 13.7pp for the highest income quintile. Two-thirds of the estimated difference in price growth experienced by households in the lowest and highest income quintiles is due to differences in log pass-through resulting from complete pass-through in levels, while one-third is due to secular differences in inflation rates.

Related literature. This paper relates to a large literature that studies theoretical and empirical determinants of pass-through.[3] I focus on the long-run pass-through of commodity shocks that shift costs for all firms in a market. Thus, I abstract from two topics that have generated large empirical literatures: (1) asymmetry in the transmission of cost increases *vs.* decreases (e.g., Borenstein et al. 1997; Peltzman 2000; Benzarti et al. 2020) and (2) the pass-through of idiosyncratic shocks that only affect some firms in a market (as in much of the literature on exchange rate pass-through, e.g., Campa and Goldberg 2005; Berman et al. 2012; Burstein and Gopinath 2014; Amiti et al. 2019).[4]

While this paper is the first to propose complete pass-through in levels as a pattern spanning several markets, there are previous studies that measure pass-through in levels in certain contexts. For example, Nakamura and Zerom (2010) find that retail coffee prices move one-for-one with coffee commodity prices. However, the central exercise in Nakamura and Zerom (2010) seeks to account for the incomplete pass-through in logs, attributing long-run incomplete log pass-through to both non-commodity input costs and a positive super-elasticity of demand. Studies of gasoline markets also typically measure pass-through in levels rather than in logs (e.g., Karrenbrock 1991; Borenstein et al. 1997; Deltas 2008). However, these studies do not provide a theoretical foundation for why complete pass-through in levels is an appropriate benchmark.[5] Other papers

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[3]See e.g., Bulow and Pfleiderer (1983), Leibtag (2009); Nakamura and Zerom (2010); Weyl and Fabinger (2013); Hong and Li (2017); Minton and Wheaton (2022).

[4]Complete pass-through in levels may explain some patterns documented in this literature, such as the fact log pass-through falls with products quality (Chen and Juvenal 2016) and firm size (e.g., Berman et al. 2012, Gupta 2020). Both patterns would result markups increase with quality and size.

[5]Borenstein (1991) notes, "Though standard economic theory indicates that the percentage markup over marginal cost is the correct measure of market power, the industry literature and analysis focuses on the

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that measure pass-through in levels include Dutta et al. (2002), who document roughly complete pass-through in levels for frozen orange juice concentrate, and Conlon and Rao (2020), who find pass-throughs in levels over one due to discrete price points in distilled spirits.

Most closely related to my study of pass-through in levels is Butters et al. (2022), who study how retail stores' prices respond to local cost shocks such as excise tax changes. Consistent with my evidence, Butters et al. (2022) find evidence of complete pass-through in levels of these cost changes.[6] This paper adds to this evidence by showing that that complete pass-through in levels is not a feature unique to retail stores, but holds along the chain of producers from commodity to retailer in the studied markets.[7]

The application to inflation inequality builds on a rich literature exploring differences in inflation across households (Hobijn and Lagakos 2005; Kaplan and Schulhofer-Wohl 2017; Jaravel 2019, 2021). My findings confirm the secular differences in inflation across income groups documented in this literature, and also point to a new source of inflation inequality that varies with upstream costs. This channel is relevant for understanding why inflation inequality can surge when commodity prices are rising.

Finally, the model of additive unit margins in this paper also relates to a large literature on so-called full cost or cost-plus pricing. This literature was spawned by survey evidence that pricing managers predominantly use simple heuristics to set prices, and often set pricing based on average costs rather than marginal costs (see e.g., Hall and Hitch 1939; Kaplan et al. 1958; Lanzillotti 1958). Historical debates between marginal cost and full cost pricing theories are surveyed in Heflebower (1955) and Okun (1981). The model of safety margin constraint pricing developed in this paper draws on the verbal discussion of full cost pricing in Fellner (1948).

Layout. Section 2 motivates the empirical specifications used to measure pass-through in logs and in levels. Section 3 documents empirical patterns of pass-through in the retail gasoline market. Section 4 examines pass-through in food product markets. Section 5 explores explanations for pass-through in levels, develops a model of firm pricing, and tests its empirical predictions. Section 6 applies pass-through in levels to the unequal incidence of commodity shocks. Section 7 concludes.

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retail/wholesale margin measured in cents." He suggests that this may be because retail gas stations' market power is derived from consumers' time cost of visiting other stations.

[6]Butters et al. (2022) note that their results "are consistent with perfect competition, [though] perfect competition is inconsistent with the substantial evidence that retailers exhibit some degree of market power." This puzzle is the central puzzle addressed in this paper.

[7]In Appendix E, I estimate pass-through at different stages of the agricultural supply chain (farm to wholesale to retail). I find evidence of complete pass-through in levels at each step of the supply chain.

2 Framework and Empirical Specification

Pass-through in logs and levels. To fix ideas, consider a firm that produces an output good using both commodity and non-commodity inputs. In particular, we will assume that the firm has a constant returns, Leontief production technology, so that the cost of producing y units of the output good is $C(y)$:[8]

$$C(y) = y(c + w),$$

where the price of the commodity input is c and the price of the non-commodity input is w (with units of each good required to produce one unit of the output good normalized to one). Table 1 shows an example in which $c = \$1$ and $w = \$1$.

In standard models where buyers have CES preferences over varieties (e.g., Dixit and Stiglitz 1977), firms' desired prices p^* are equal to a fixed, multiplicative markup, μ, over marginal cost:

$$p^* = \mu(c + w). \tag{1}$$

In Table 1, this markup is initially $\mu = 2$, resulting in an initial output price of $2(\$1+\$1) = \$4$.

How does an increase in the commodity price, Δc, affect the price charged by the firm? Under the multiplicative pricing rule in (1), the change in the firm's desired price is

$$\Delta p^* = \mu \Delta c.$$

Thus, when a firm sets a fixed multiplicative markup over cost, the pass-through in levels of commodity price changes is equal to the markup μ. In markets with imperfect competition, $\mu > 1$, and hence the standard model with fixed multiplicative markups implies pass-through in levels greater than one.

Table 1 row (a) shows the pass-through that results from a fixed multiplicative markup rule for a \$0.20 increase in the commodity price. Since a \$0.20 increase in the commodity price increases marginal costs by 10 percent, complete log pass-through implies that the output price should also rise by 10 percent, or by \$0.40. The pass-through in levels—the ratio of the change in the output price to the change in the commodity cost—is equal to the markup, $\mu = 2$.[9]

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[8]The assumption of constant returns, Leontief production seems appropriate for the settings that I study empirically. In order to sell an ounce of ground coffee, a firm must buy the equivalent amount of coffee beans. Section 5 considers how relaxing the assumptions of constant returns to scale, Leontief production, and isoelastic demand each affect pass-through.

[9]If pass-through in logs is calculated as the percentage change in the output price relative to the commodity cost, rather than marginal cost, the resulting log pass-through is equal to the cost share: 10%/20% = 0.5.

Table 1: Example of pass-through in logs and levels.

	Start		New	% Change	Pass-through	
					Logs	Levels
Components of marginal cost						
Commodity	$1	+$0.20	$1.20	+20%		
Other variable costs	$1	–	$1.00			
Total marginal cost	$2	+$0.20	$2.20	+10%		
Desired output price						
(a) Fixed multiplicative markup	$4	+$0.40	$4.40	+10%	1.0	2.0
(b) Fixed additive margin	$4	+$0.20	$4.20	+5%	0.5	1.0

Suppose that the firms' desired price is instead described by the additive pricing rule

$$p^* = c + w + \alpha,$$

where α is a fixed, unit margin between the output price and marginal cost. In Table 1, the initial $\alpha = \$2$.

Under (2), an increase in the commodity price results in a one-for-one increase in the output price (row (b)). When measured on a percentage basis, the change in the output price (5 percent) appears incomplete relative to the change in marginal cost (10 percent). The percent change in the output price relative to the commodity price (5 percent / 20 percent = 0.25) is also incomplete relative to the commodity's initial cost share (0.5). In other words, complete pass-through in levels is disguised as incomplete log pass-through.

Empirical specification. The aim of the empirical strategy is to measure how firms' desired prices p^* respond to changes in commodity costs, both in logs and in levels. Of course, at short horizons, price rigidities may prevent a firm from setting its posted price p in accordance with changes in its desired price p^*. Hence, I study pass-through of persistent cost changes at long horizons where price rigidities are overcome.

I estimate this long-run pass-through of commodity cost changes using the standard distributed lag regression (see e.g., Campa and Goldberg 2005, Nakamura and Zerom 2010),

$$\Delta p_t = a + \sum_{k=0}^{K} b_k \Delta c_{t-k} + \epsilon_t, \tag{2}$$

where Δp_t is the change in the output price (in levels) from $t-1$ to t, Δc_{t-k} is the change in

the commodity cost (in levels) from $t-k-1$ to $t-k$, and ϵ_t is a mean zero error term.[10]

The estimated coefficients b_k measure the change in the output price associated with a change in commodity costs k periods ago. Accordingly, the long-run pass-through of a change in the commodity cost Δc to prices is given by the sum of the coefficients, $\sum_{k=0}^{K} b_k$.[11]

I estimate the long-run "log pass-through" using the analogous specification,

$$\Delta \log p_t = \alpha + \sum_{k=0}^{K} \beta_k \Delta \log c_{t-k} + \epsilon_t, \tag{3}$$

where the long-run log pass-through is $\sum_{k=0}^{K} \beta_k$.

For valid inference, changes in the commodity cost Δc_t must be stationary. I confirm this is the case for all commodity series used in this paper in Appendix Table A1. In all cases, autocorrelation estimates and Augmented Dickey-Fuller tests suggest that while the commodity costs series are approximately unit root, first differences in commodity costs are not.

3   Evidence from Retail Gasoline

Retail gasoline provides an ideal laboratory to study pass-through since rich data on upstream costs is readily available and gasoline prices exhibit little rigidity. My main analysis in this section uses data on the universe of retail gas stations in Perth, Australia. At the end of the section, I show that retail gasoline markets in the United States, Canada, and South Korea exhibit similar patterns.

I document four patterns in the retail gasoline data. First, estimates of the pass-through in levels from wholesale prices to retail prices are statistically indistinguishable from one. Second, long-run log pass-through is incomplete even relative to the share of gasoline in stations' marginal costs.[12] Third, there is little heterogeneity in pass-through in levels across stations in the sample, but substantial variation in log pass-through: stations with a larger gap between prices and costs have lower log pass-through. Using several instruments designed to isolate variation in stations' markups from stations' marginal costs, I find that stations with higher markups have lower log pass-through. Fourth,

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[10]An alternative is to use a vector error correction (VEC) model, which allows for co-integrated cost and price series. Using these specifications produces broadly similar results to my baseline results; however, the estimates are substantially noisier.

[11]One can show that for any time-dependent price rigidities, when c is a unit root process, under the constant markup pricing rule (1), the long-run pass-through estimate $\sum_{k=0}^{K} b_k \to \mu$ as K becomes large.

[12]Weekly price cycles in the data (documented previously by Wang 2009a and Byrne and de Roos 2019) allow me to estimate a lower bound on the marginal cost share of gasoline.

complete pass-through in levels and variation in stations' margins explain both cross-sectional heterogeneity in log pass-through and the overall level of incomplete log pass-through.

3.1 Station-Level Data from Perth, Australia

Station-level retail gasoline price data are from FuelWatch, a Western Australia government program that has monitored retail gasoline prices since January 2001. Alongside the introduction of the FuelWatch program in 2001, the Western Australian government banned intra-day price changes and required all retail gas stations to submit prices for each gas product (i.e., unleaded petrol, premium unleaded petrol, and diesel) by 2pm of the prior day. Since 2003, FuelWatch also provides daily data on the local spot price for wholesale gasoline, called the terminal gas price, across six terminals used by retail stations. (Previous studies using these data include Wang 2009a; Byrne and de Roos 2017, 2019, 2022).

Following Byrne and de Roos (2019), I take the minimum terminal gas price offered by the six terminals each day as the commodity price that faces retail gas stations. Figure 2 shows the weekly average terminal gas price and the retail unleaded petrol (ULP) price for a single gas station from 2001 to 2022. The retail price is slightly above, but closely tracks, the terminal gas price. The gap between retail and wholesale prices visibly increases in 2010. Byrne and de Roos (2019) document that retail gas margins in Perth increased starting in 2010 due to the emergence of tacit collusion across stations, a feature of the market that I exploit later in the analysis.[13]

3.2 Empirical Results

Pass-through is complete in levels and incomplete in logs. Figure 3 shows the estimated pass-through of changes in unleaded petrol (ULP) wholesale prices to station retail prices over a horizon of eight weeks. By three weeks, the pass-through in levels is statistically indistinguishable from one. The point estimate for long-run pass-through at eight weeks is 0.991 (standard error 0.038). Changing the horizon over which pass-through is estimated has little effect on the estimated long-run pass-through. In contrast, the log pass-through at eight weeks is 0.899 (0.043) and is statistically different from one at a 1 percent level.

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[13]While the BP station in Kewdale shown in Figure 2 tracks wholesale prices quite closely, as is typical of most gas stations in the sample, some gas stations maintain prices significantly higher than the wholesale price and update prices less frequently. See, for example, ULP prices for the Rottnest Island Authority station shown in Appendix Figure A1.

Figure 2: Weekly average retail unleaded petrol (ULP) price and terminal gas price for BP station at 549 Abernethy Rd, Kewdale, Perth, Australia.
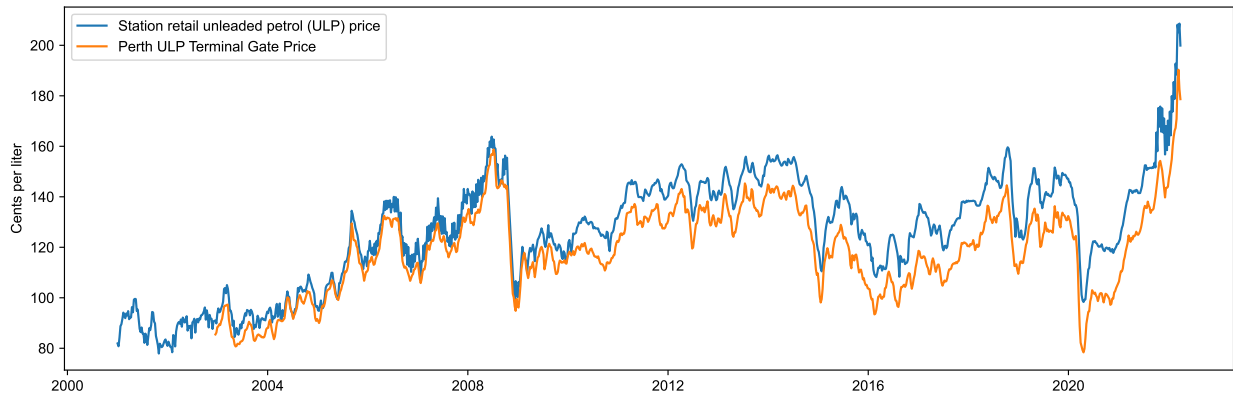

Estimates of the pass-through of premium unleaded (PULP) wholesale prices to retail prices (Appendix Figure A2) are similar: the long-run pass-through in levels is statistically indistinguishable from one at 0.985 (0.036), while the long-run pass-through in logs is significantly below one at 0.887 (0.041).

Incomplete log pass-through even accounting for cost share. If retail gas stations face other variable costs besides the cost of gasoline, and stations use a multiplicative markup pricing rule, the log pass-through should be equal to the share of gasoline in stations' marginal costs. Luckily, the presence of price cycles in this setting allows us to estimate a lower bound for the cost share of gasoline and observe whether log pass-through is complete after accounting for the gasoline cost share.

Figure 4 shows daily prices charged by a single gas station in the sample from March to June 2016. As previously documented by Byrne and de Roos (2019), the retail price follows weekly price cycles, typically jumping up on Tuesdays and then falling over the course of the week. Under the assumption that gas stations never set prices below marginal cost,[14] we can use the days of the week at the lowest point of the price cycle to calculate an upper bound on the share of other variable costs in stations' marginal costs, and thus a lower bound for the cost share of gasoline. This approach yields a lower bound for the cost share of gasoline of 0.98 for unleaded petrol and 0.96 for premium unleaded petrol. The estimated log pass-throughs for unleaded and premium unleaded petrol, at 0.899 and 0.887, are significantly different from these cost shares at the 1 percent level. Thus, the estimated log pass-through of gasoline costs is incomplete, even accounting for the share

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[14]This is the case in the seminal Maskin and Tirole (1988) model of price cycles.

Figure 3: Unleaded petrol (ULP) price pass-through in levels (top) and in logs (bottom).
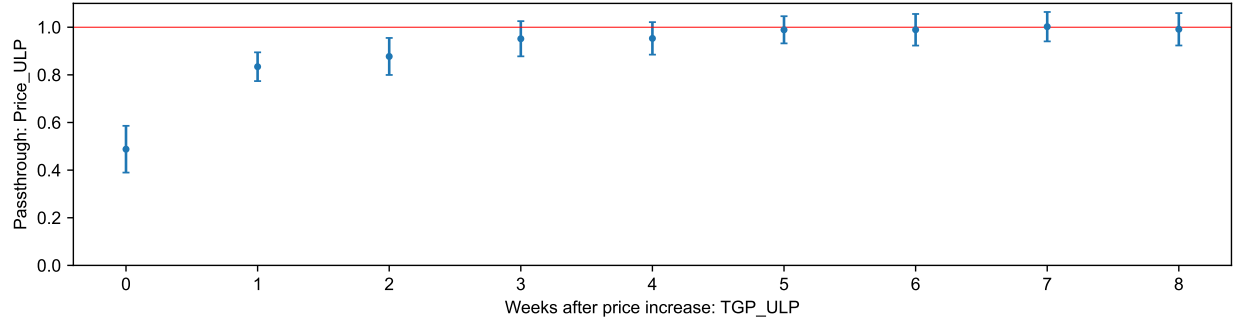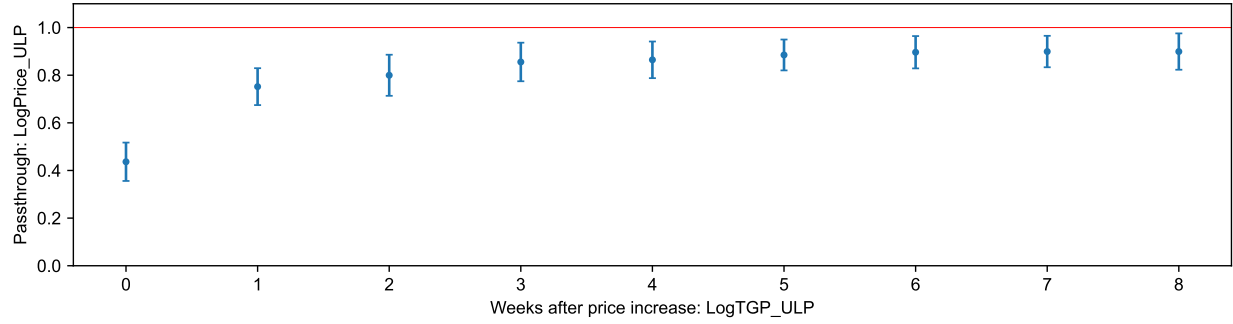
(a) Pass-through in levels.

(b) Pass-through in logs.

Note: Panels (a) and (b) show cumulative pass-through estimated from the specifications,

$$\Delta p_{i,t} = \sum_{k=0}^{k=8} b_k \Delta c_{i,t-k} + a_i + \varepsilon_{i,t}.$$

$$\Delta \log p_{i,t} = \sum_{k=0}^{k=8} \beta_k \Delta \log c_{i,t-k} + \alpha_i + \varepsilon_{i,t}.$$

Standard errors are two-way clustered by postcode and year (Driscoll-Kraay panel standard errors are similar), and standard errors for cumulative pass-through coefficients $\sum_{k=0}^{t} b_k$ and $\sum_{k=0}^{t} \beta_k$ are computed using the delta method.

Figure 4: Daily retail unleaded petrol price for BP at 549 Abernethy Rd, Kewdale, Perth for three months in 2016, with lowest points in price cycle.
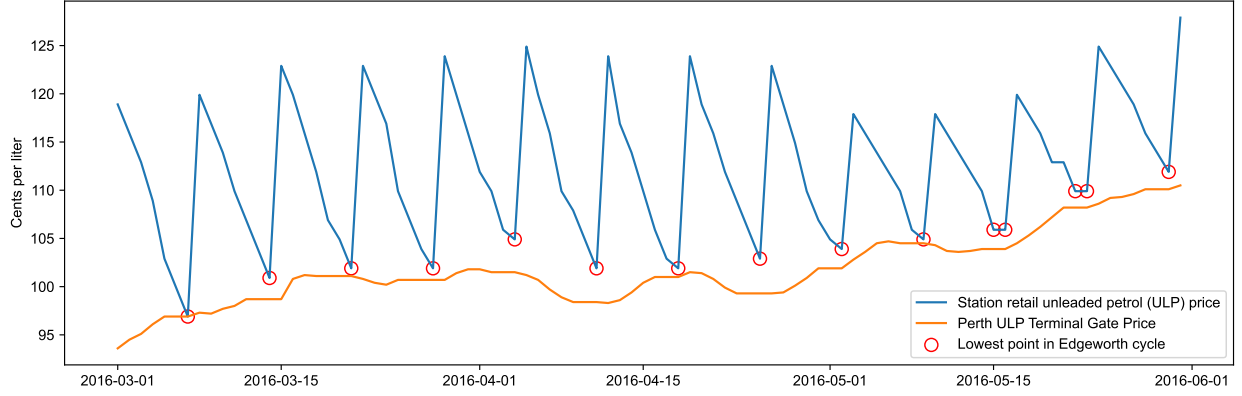
of gasoline in variable costs.[15]

Exploiting variation in markups. While the point estimate for pass-through in levels (0.991) is very close to one, it is hard to reject low, single-digit multiplicative markups that would be plausible in this setting. To investigate this, I exploit cross-sectional and time series variation in markups. If stations follow a multiplicative markup pricing rule, and if some stations charge higher markups than others, then the pass-through in levels for high-markup stations should be higher than their low-markup counterparts.

I estimate the specification,

$$\Delta p_{i,t} = \alpha + \delta \Delta c_t + \gamma \text{Avg. Markup}_{i,t} + \beta(\Delta c_t \times \text{Avg. Markup}_{i,t}) + \varepsilon_{i,t}.$$

where $\Delta p_{i,t}$ and Δc_t are changes in the station retail price and wholesale cost over the prior sixteen weeks, Avg. Markup$_{i,t}$ is a measure that exploits cross-sectional or time series variation in markups, and $\varepsilon_{i,t}$ is a mean-zero error term.[16]

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[15]One may wonder whether the fact that log pass-through is lower than the average cost share could be due to higher order terms. If anything, higher order terms should upwardly bias log pass-through relative to the cost share. To see why, suppose stations were perfectly competitive ($p = c + w$), so that log pass-through to a first-order were equal to the cost share ($\chi = c/(c + w)$). The change in log prices to a second order is

$$\Delta \log p \approx \chi(d \log c) + \chi(1 - \chi)(d \log c)^2.$$

The estimate of log pass-through will be

$$\hat{\rho} = \mathbb{E}[\Delta \log p / d \log c] \approx \mathbb{E}[\chi] + \mathbb{E}[\chi(1 - \chi)(d \log c)].$$

The second order term makes log pass-through higher than the cost share if the commodity price has upward drift ($\mathbb{E}[d \log c] > 0$).

[16]Since I find that pass-through is complete over an eight-week horizon, I take changes in station retail

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The constant markup model predicts that the coefficient on the interaction term $\beta > 0$. For example, if some stations set a constant 2 percent markup and other stations set a constant 5 percent markup, pass-through in levels should be 1.05 for the high-markup stations compared to 1.02 for the low-markup stations (thus, $\beta > 0$). On the other hand, if all stations use an additive margin pricing rule and obey complete pass-through in levels, the interaction coefficient $\beta \approx 0$.

I use two proxies for Avg. Markup$_{i,t}$ along with instruments for both. The first measure exploits variation in markups across stations: Avg. Station Markup$_i$ is the average markup (gasoline retail price / wholesale price) charged by station i over all weeks in the sample. I also instrument for Avg. Station Markup$_i$ with the average amplitude of price cycles of station i, that is, the average difference between the maximum and minimum retail margin charged by i in each week. While the price / cost measure of a station's markup may also capture variation in non-gasoline variable costs, this instrument measures markups using only the gap between a station's maximum and minimum margin.

The second measure instead exploits variation in markups over time: Avg. Markup$_t$ is the average markup across all gas stations in year t. To instrument for Avg. Markup$_t$, I take advantage of the fact that the emergence of coordinated price cycles in the Perth market was, according to Byrne and de Roos (2019), "unrelated to market primitives." Over the course of the sample, average gas station margins co-move closely with the degree of coordination in price cycles, measured as the R^2 from a regression of daily margins on day-of-week fixed effects (see Appendix Figure A3).[17] I use this measure of price coordination over time—the quarterly R^2 of station margins on day-of-week dummies—as an instrument for the average year margin Avg. Margin$_t$.

Table 2 reports the results. In column 1, I omit the average markup and interaction term. A \$1 change in the wholesale cost of unleaded petrol (ULP) over 16 weeks is associated with a \$0.95 change in the retail station price over 16 weeks.[18] Columns 2–5 include the interaction of the cost change with the average station or annual markup, both reduced form and instrumenting for the markup using the strategies discussed above. In all cases, the $\beta \approx 0$, rejecting the prediction of the constant-markup model. In other

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prices and costs over twice this period, or sixteen weeks (i.e., $\Delta p_{i,t} = p_{i,t} - p_{i,t-16}$). At this horizon, cost changes in the first eight weeks should be fully passed through. Similar results obtain using different lengths choices of the price change horizon.

[17]While the most dramatic change over time is the increase in both coordination and margins around 2010, there is also subsequent variation in the strength of coordination and margins after 2010, which may owe to subsequent price wars documented by Byrne and de Roos (2019).

[18]Since pass-through is only complete at horizons of 5–8 weeks, changes in cost in the final weeks of this 16 week difference, $\Delta c_t = c_t - c_{t-16}$, will not be completely passed through to prices, and hence this coefficient is slightly less than one.

Table 2: Complete pass-through in levels: No heterogeneity by station markup.

ΔPrice_{it}	(1) (OLS)	(2) (OLS)	(3) (IV1)	(4) (OLS)	(5) (IV2)
ΔCost_t	0.950**	0.989**	0.952**	0.985**	0.973**
	(0.021)	(0.037)	(0.044)	(0.043)	(0.048)
$\Delta\text{Cost}_t \times$ Avg. Station Markup$_i$ (Net %)		-0.005	-0.000		
		(0.003)	(0.005)		
$\Delta\text{Cost}_t \times$ Avg. Year Markup$_t$ (Net %)				-0.004	-0.003
				(0.004)	(0.005)
N	312215	312215	312215	312215	312215
R^2	0.89	0.89	0.89	0.89	0.89

Note: The table reports the coefficients γ and β from the specification,

$$\Delta\text{Price}_{i,t} = \alpha + \delta\Delta\text{Cost}_t + \gamma\text{Avg. Markup}_{i,t} + \beta(\Delta\text{Cost}_t \times \text{Avg. Markup}_{i,t}) + \varepsilon_{i,t}.$$

Changes in prices and costs ΔPrice_{it} and ΔCost_t are taken over 16 weeks, and Avg. Markup$_{i,t}$ is included on a net % basis (i.e., a markup of $p/c = 1.1$ is input as 10%). Column 3 (IV1) uses station i's average price cycle amplitude as an instrument for Avg. Margin$_i$. Column 5 (IV2) uses the quarterly R^2 of station margins on day-of-week dummies as an instrument for Avg. Margin$_t$. Standard errors are two-way clustered by postcode and year.

words, there are no detectable deviations from complete pass-through in levels in the cross-section of stations or in the time series.[19]

Pass-through in levels explains heterogeneity in log pass-through. Table 3 reports estimates from an analogous specification that instead measures the pass-through of changes in log costs to changes in log prices,[20]

$$\Delta\log p_{i,t} = \alpha + \beta\Delta\log c_{i,t} + \delta\text{Avg. Markup}_{i,t} + \gamma(\Delta\log c_{i,t} \times \text{Avg. Markup}_{i,t}) + \varepsilon_{i,t}. \quad (4)$$

Column 1 omits the average markup and interaction term and estimates an average log pass-through from cost changes to price changes of 0.870. Thus, like the long-run pass-through estimates, this log pass-through is significantly lower than one. Columns

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[19]As a robustness check, I compare the long-run pass-through for stations grouped by their price relative to other stations in the same neighborhood. Assuming unobserved costs (e.g., transport, rent) are the same within a neighborhood, stations with higher relative prices have higher markups and thus should exhibit higher pass-through in levels if prices are set with fixed multiplicative markups. Appendix Figure A4 shows that long-run pass-through in levels is close to one and unchanging across all groups of relative prices within neighborhoods and within postcodes.

[20]Since Table 2 suggests that stations set a constant additive margin, rather than a multiplicative markup, one might find it preferable to estimate specification (4) using an interaction with Avg. Margin$_{i,t}$ rather than Avg. Markup$_{i,t}$. Results from that specification are quantitatively similar to the results in Table 3.

Table 3: Incomplete log pass-through is explained by station margins.

$\Delta \log(\text{Price})_{it}$	(1) (OLS)	(2) (OLS)	(3) (IV1)	(4) (OLS)	(5) (IV2)
$\Delta \log(\text{Cost})_t$	0.870**	0.998**	0.968**	0.985**	0.977**
	(0.031)	(0.035)	(0.041)	(0.035)	(0.040)
$\Delta \log(\text{Cost})_t \times$ Avg. Station Markup$_i$ (Net %)		-0.015**	-0.011**		
		(0.003)	(0.004)		
$\Delta \log(\text{Cost})_t \times$ Avg. Year Markup$_t$ (Net %)				-0.012**	-0.011**
				(0.003)	(0.004)
N	312215	312215	312215	312215	312215
R^2	0.88	0.89	0.89	0.89	0.89

Note: The table reports the coefficients γ and β from the specification,

$$\Delta \log(\text{Price})_{i,t} = \alpha + \delta \Delta \log(\text{Cost})_t + \gamma \text{Avg. Markup}_{i,t} + \beta(\Delta \log(\text{Cost})_t \times \text{Avg. Markup}_{i,t}) + \varepsilon_{i,t}.$$

Changes in log prices and costs $\Delta \log(\text{Price})_{it}$ and $\Delta \log(\text{Cost})_t$ are taken over 16 weeks, and Avg. Markup$_{i,t}$ is included on a net % basis (i.e., a markup of $p/c = 1.1$ is input as 10%). Column 3 (IV1) uses station i's average price cycle amplitude as an instrument for Avg. Margin$_i$. Column 5 (IV2) uses the quarterly R^2 of station margins on day-of-week dummies as an instrument for Avg. Margin$_t$. Standard errors are two-way clustered by postcode and year.

2–5 include the average markup and interaction term, again exploiting cross-sectional variation in markups (columns 2–3) or time series variation in markups (columns 4–5). Two findings emerge. First, higher margins (and thus higher markups) lead to a measured log pass-through that is more incomplete.[21] Second, the gap between price and costs appears to fully account for incomplete pass-through: the coefficient on $\Delta \log(\text{Cost})_t$ shows that as the net markup approaches zero, the log pass-through is tightly estimated around the cost share of 0.98.

Thus, Table 3 shows that incomplete log pass-through is rationalized by the combination of complete pass-through in levels (documented in Table 2) with non-zero margins. In particular, because pricing follows complete pass-through in levels, log pass-through is lower both for stations in the cross-section with higher margins and for years in the time-series with higher margins. The presence of additive unit margins between costs and prices explains both the level of incomplete log pass-through (around 0.87) and variation

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[21]The estimated interaction coefficient in the log specification $\beta \approx -0.01$ is not a coincidence. If stations set an additive unit margin α over marginal cost $c + w$, to a first order,

$$\Delta \log p \approx \frac{c}{c + w + \alpha} \Delta \log c = \chi \mu^{-1} \Delta \log c \approx \chi(1 - 0.01\mu^{\text{net},\%})\Delta \log c,$$

where $\chi = c/(c+w)$ is the cost share (0.96–0.98 in the data), $\mu = p/(c+w)$ is the markup, and $\mu^{\text{net},\%} = 100(\mu-1)$.

in log pass-through across stations.

Evidence from other markets and oil supply shocks. Table 4 compares pass-through estimates from Perth to estimates from a panel of Canadian cities and from station-level data from South Korea (Appendix C describes the data sources used to construct these estimates). Incomplete log pass-through and complete pass-through in levels appear across all the studied markets.[22] Appendix Figure A6 shows that complete pass-through in levels also appears in U.S. data, using refiner wholesale prices from the Energy Information Administration (EIA) and retail prices from the BLS. The evidence from other geographies suggests that complete pass-through in levels is not a quirk of the Australian data, but rather describes price dynamics across a number of retail gasoline markets.

Table 4: Pass-through estimates: Other geographies and Känzig (2021) instrument.

| | Long-run pass-through (8 weeks) | | | |
| | Logs | | Levels | |
Description	Est.	IV	Est.	IV
Australia, station-level, 2001–2022				
Terminal to retail, Unleaded	0.899	0.805[†]	0.991[†]	0.888[†]
	(0.043)	(0.118)	(0.038)	(0.132)
Terminal to retail, Premium Unleaded	0.887	0.812[†]	0.985[†]	0.901[†]
	(0.041)	(0.129)	(0.036)	(0.146)
Canada, city-level, 2007–2022				
Crude to wholesale	0.553	0.713	0.927[†]	1.086[†]
	(0.098)	(0.146)	(0.100)	(0.186)
Wholesale to retail (excl. taxes)	0.859	0.848	1.008[†]	0.994[†]
	(0.016)	(0.042)	(0.022)	(0.049)
South Korea, station-level, 2008–2022				
Refinery to retail, Unleaded	0.926[†]	0.935[†]	0.997[†]	1.012[†]
	(0.044)	(0.097)	(0.052)	(0.108)

Note: The table reports long-run pass-through at a horizon of eight weeks for station-level data from Perth, Australia, city-level data from Canada, and station-level data from South Korea. Driscoll-Kraay standard errors with eight lags in parentheses. The IV columns use eight lags of OPEC announcement shocks from Känzig (2021) as an instrument for commodity price changes. The *F*-stat for the instrument in all regressions is greater than 10. [†] indicates that an estimate is statistically indistinguishable from one.

So far, we have assumed that commodity costs pass downstream to retail prices and

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[22]For Canada, city-level wholesale prices are also available, allowing me to estimate pass-through from crude to city-level wholesale prices. Interestingly, these data also suggest complete pass-through in levels from crude to wholesale prices.

not vice versa (this assumption is supported by the Granger causality tests in Appendix Table A2). As an additional check, Table 4 also estimates pass-through instrumenting for upstream commodity cost changes with OPEC announcement shocks from Känzig (2021). Estimates of the long-run pass-through in levels and logs from the instrumented regressions are similar to the baseline results.

4 Evidence from Food Products

To investigate whether these empirical patterns hold in other markets beyond retail gasoline, in this section I explore pass-through of commodity costs to retail prices in six staple food products (coffee, sugar, ground beef, white rice, all-purpose flour, and frozen orange juice concentrate).

For five out of the six product, the long-run pass-through of commodity costs in levels is statistically indistinguishable from one.[23] Using product-level scanner data for three food products (coffee, rice, and flour), I find that products in the cross-section with higher unit prices have lower log pass-through, but have no systematic differences from low unit-price products in pass-through in levels. Like in the cross-section of retail gasoline stations, variation in log pass-through across products in a category can be rationalized by variation in non-commodity marginal costs and margins.

Finally, I document that these patterns in pass-through appear to extend to a broader set of fast-moving goods, by exploiting the fact that different retailers often set different prices for identical products (Kaplan and Menzio 2015; Kaplan et al. 2019). The behavior of prices of identical products across retailers conform to the predictions of an additive unit margin pricing model, rather than a fixed multiplicative markup model.

4.1 Data on Food Retail and Commodity Prices

Retail prices. For retail prices of food products, I primarily rely on Average Price Data from the Bureau of Labor Statistics. While most BLS CPI series capture relative price changes, the Average Price Data track price levels for a select number of staple products. For each price series, the BLS chooses narrowly defined, homogeneous item categories to minimize input or quality differences between included items. In many cases, Average Price Data also defines items by package size (e.g., "Orange juice, frozen concentrate, 12 oz. can, per 16 oz."), thereby minimizing differences in size that could affect price levels.

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[23]The exception is sugar, where the estimate of long-run pass-through in levels is 0.691.

While the BLS Average Price Data allow us to study pass-through of commodity costs to retail prices over a long time series—many of the series record prices back to 1980—studying cross-sectional heterogeneity across products in a category requires richer data. For these investigations, I use Nielsen Retail Scanner data, which provides weekly barcode-level prices and quantities for products sold at participating stores from 2006 to 2020. These data are collected from point-of-sale systems in about 90 retail chains operating across the U.S., reflecting over $2 billion in annual sales.

Commodity costs. I match retail food prices with data on commodity costs from the IMF Primary Commodities Prices database. These commodity price series draw from statistics of specialized trade organizations or from commodity futures markets—for example, the U.S. sugar commodity price from the IMF uses the price of the nearest Sugar No. 16 futures contract, which is for delivery of cane sugar from the U.S. or another duty-free origin to New York, Baltimore, Galveston, New Orleans, or Savannah. Appendix Table A3 provides a full list of the commodity price series used and the underlying data sources used by the IMF.

Measuring pass-through in levels requires carefully matching units from commodity prices to retail prices. For example, to measure pass-through of wheat commodity prices to retail flour prices requires knowing the quantity of wheat needed per pound of flour produced. To construct these mappings from commodity units to retail units, I rely on previous literature and on sources provided by the USDA. Appendix Table A4 provides the conversion factors from commodity prices to retail prices for each series and delineates the sources and assumptions used to build each conversion factor.[24]

Matched products. Of the food products tracked by the BLS Average Price Data, six can be clearly matched to commodity input prices provided by the IMF. These are roasted ground coffee, sugar, ground beef, white rice, all-purpose flour, and frozen orange juice concentrate. Appendix A4 lists the corresponding Average Price Data Series IDs and reported units. For three of these products—rice, flour, and coffee—I also investigate cross-sectional pass-through patterns by matching the food product to a Nielsen product category.[25]

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[24]This careful matching of units is one of the reasons why estimating pass-through in levels is difficult for non-staple, differentiated products, where the portion of costs attributed to upstream commodity prices is more difficult to assess than for these homogeneous, staple goods. At the end of this section, I show that several other products exhibit similar pass-through patterns to the homogeneous goods studied in this section, but I do so without the ability to match retail products directly to their commodity inputs.

[25]The corresponding Nielsen product modules are Nielsen product module 1319 "Rice - Packaged and bulk" for rice, Nielsen product module 1393 "Flour - All purpose - White wheat" for flour, and Nielsen

Table 5: Long-run pass-through of commodity costs to retail food prices.

Commodity (IMF)	Final Good (BLS)	Logs		Levels	
Arabica coffee	Coffee, 100%, ground roast	0.466	(0.051)	0.946[†]	(0.099)
Sugar, No. 16	Sugar, white	0.370	(0.035)	0.691	(0.072)
Beef	Ground beef, 100% beef	0.410	(0.068)	0.899[†]	(0.126)
Rice, Thailand	Rice, white, long grain, uncooked	0.307	(0.049)	0.882[†]	(0.169)
Wheat	Flour, white, all purpose	0.240	(0.048)	0.819[†]	(0.152)
Frozen orange juice	Orange juice, frozen concentrate	0.327	(0.040)	1.006[†]	(0.114)

Note: Long-run pass-through in levels and logs is $\sum_{k=0}^{K} b_k$ from specifications (2) and (3), using a horizon of $K = 12$ months. Newey-West standard errors in parentheses. [†] indicates that an estimate is statistically indistinguishable from one.

4.2 Empirical Results

Nearly all products exhibit complete pass-through in levels.   Table 5 reports estimates of long-run pass-through in levels and logs (specifications (2) and (3)) for six food products. In five of the six products, long-run pass-through in levels is statistically indistinguishable from one. The exception is sugar, where the estimated pass-through in levels falls short of one. For all six products, the log pass-through is significantly below one and therefore incomplete.

Figure 5 shows an example of the price series and pass-through estimates for one of the studied food products, roasted ground coffee. As shown in panel (a), Arabica coffee commodity prices exhibit substantial volatility over the period since 1980, with large spikes in 1986, 1994, 1997, 2011, and 2014 due largely to weather conditions in Brazil and Colombia.[26] These run-ups in commodity prices are followed by increases in the retail prices tracked by the BLS. Accordingly, panel (b) shows the pass-through in levels from coffee commodity prices to retail prices occurs with lags, but approaches complete pass-through by eight months and stays around one thereafter. The log pass-through, in panel (c), plateaus around 0.5. Analogous figures for the other five food products are in Appendix A.

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product module 1463 "Ground and whole bean coffee" for coffee. I exclude beef, sugar, and frozen orange juice concentrate, because beef products are spread across a number of product modules, while the "Sugar - granulated" and "Fruit juice - orange - frozen" product modules have fewer unique products.

[26]For coverage of the weather conditions leading to these coffee price run-ups, see Washington Post: "*Big Rise Predicted in Coffee Prices*" (1986), New York Times: "*Coffee Futures Soar 25%, Biggest Daily Rise in 7 Years*" (1994), New York Times: "*Coffee Hits a 20-Year High on Rumblings of a Shortage*" (1997), New York Times: "*Heat Damages Colombia Coffee, Raising Prices*" (2011), and Business Insider: "*Why Coffee Prices are Exploding*" (2014).
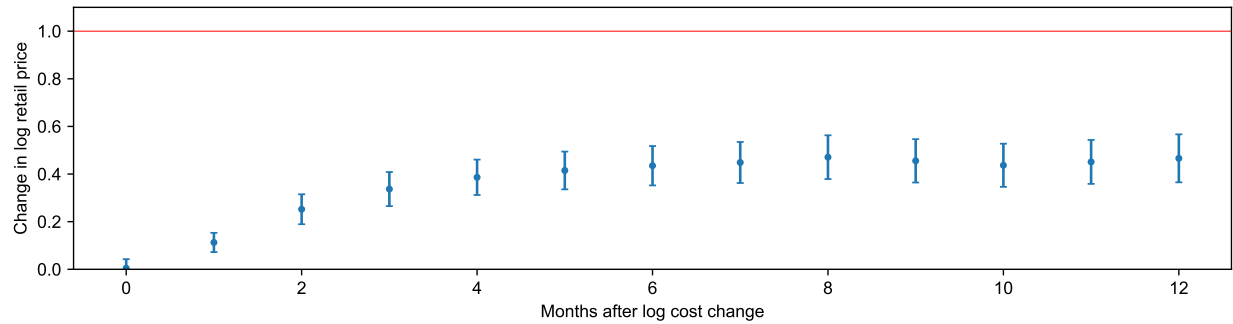
Figure 5: Passthrough of coffee commodity costs to retail prices.
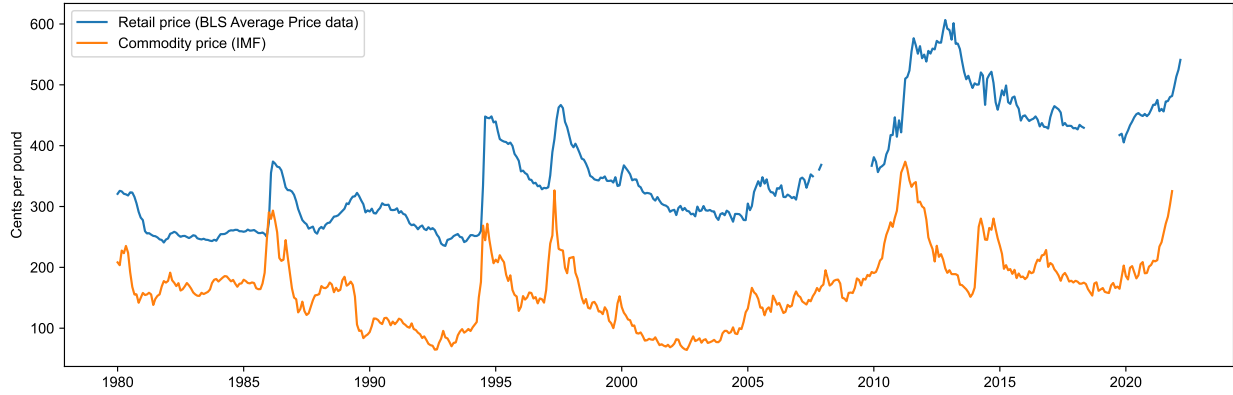
(a) Arabica coffee commodity costs (IMF) and retail ground coffee prices (U.S. CPI).
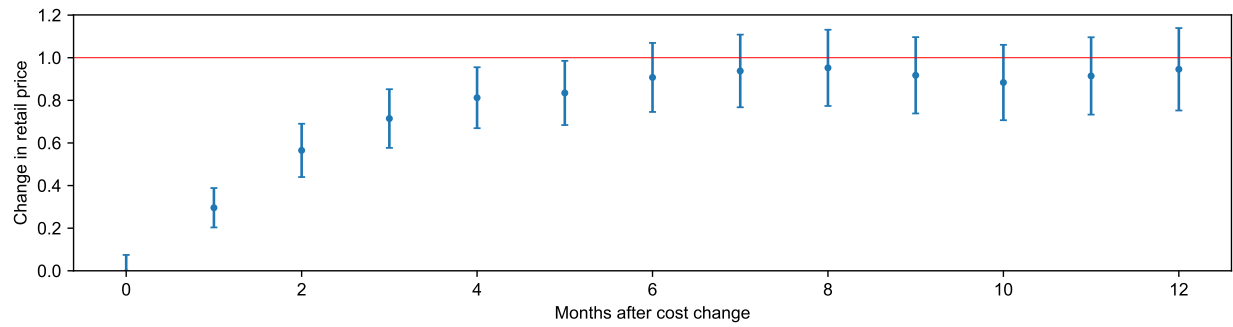
(b) Pass-through in levels.

(c) Pass-through in logs.

Note: Panel (a) plots the time series of the commodity price from the IMF and the Average Price Data series from the BLS. The series are adjusted by the conversion factors in Appendix Table A4 so that the two series are in comparable units. Panels (b) and (c) plot the cumulative pass-through to month T, $\sum_{k=0}^{T} b_k$, from the specifications (2) and (3), using a total horizon of $K = 12$ months.

Pass-through in levels explains cross-sectional variation in log pass-through. The complete pass-through in levels documented in Table 5 has predictions for price changes in the cross-section of products. First, products that have higher margins and higher non-commodity variable costs should exhibit lower log pass-through (as we saw in the cross-section of retail gas stations in Section 3). Second, pass-through in levels should be similar across products regardless of their margins and non-commodity variable costs.

To test these predictions, I use Nielsen data on rice, flour, and coffee products from 2006 to 2020. I label each unique combination of retail chain and UPC (universal product code, or product barcode) as a product.[27] In each quarter t, I calculate the price $p_{i,t}$ of retailer-UPC product i as the quantity-weighted average unit price over all transactions k,

$$p_{i,t} = \frac{\sum_k p_{i,t,k} q_{i,t,k}}{\sum_k q_{i,t,k}}.$$

For each product in each quarter, I then measure the change in the product's price over the next year in levels ($\Delta p_{i,t}$) and in percentages ($\pi_{i,t}$) as

$$\Delta p_{i,t} = p_{i,t+4} - p_{i,t}, \qquad \pi_{i,t} = (p_{i,t+4}/p_{i,t}) - 1.$$

Since these price changes are measured year over year, they avoid seasonality effects that may bias measures of price changes calculated over smaller time increments.[28]

I use the unit price (i.e., the price per ounce of rice or pound of flour) as a proxy for the extent of non-commodity variable costs and margins in the product's price. Thus, to test the above predictions for how pass-through in logs and levels varies with the level of non-commodity variable costs and margins, I group products in each product category by unit price. To ensure that these product groups capture persistent differences in unit price, I sort products by average unit price over the prior year.[29]

As an example, Figure 6 plots average inflation rates and price changes in levels for

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[27] As documented by DellaVigna and Gentzkow (2019), stores within a retail chain (especially in the same geographic area) tend to set uniform prices for each UPC at each point in time.

[28] Nakamura and Steinsson (2012) point out that using product-level data to measure pass-through may bias measurement when there is frequent product turnover. Turnover is relatively low: for these categories, 75–85 percent of quarterly retailer-UPC pairs are observed in the same quarter of the following year. Moreover, product turnover does not appear correlated with commodity inflation rates in a way that would downward bias measured pass-through. The correlation of commodity inflation over the next year with the share of products in a quarter that are also observed in the following year is 0.03 for rice products, 0.09 for flour products, and 0.09 for coffee products.

[29] Using the average unit price of the product over the prior year helps alleviate concerns that these groupings pick up temporary deviations in prices—mean reversion may then mechanically bias future price changes—though in practice all empirical results reported are very similar if we instead sort on products' unit price in only quarter t.

Figure 6: Inflation and price changes of rice products by tercile of unit price.
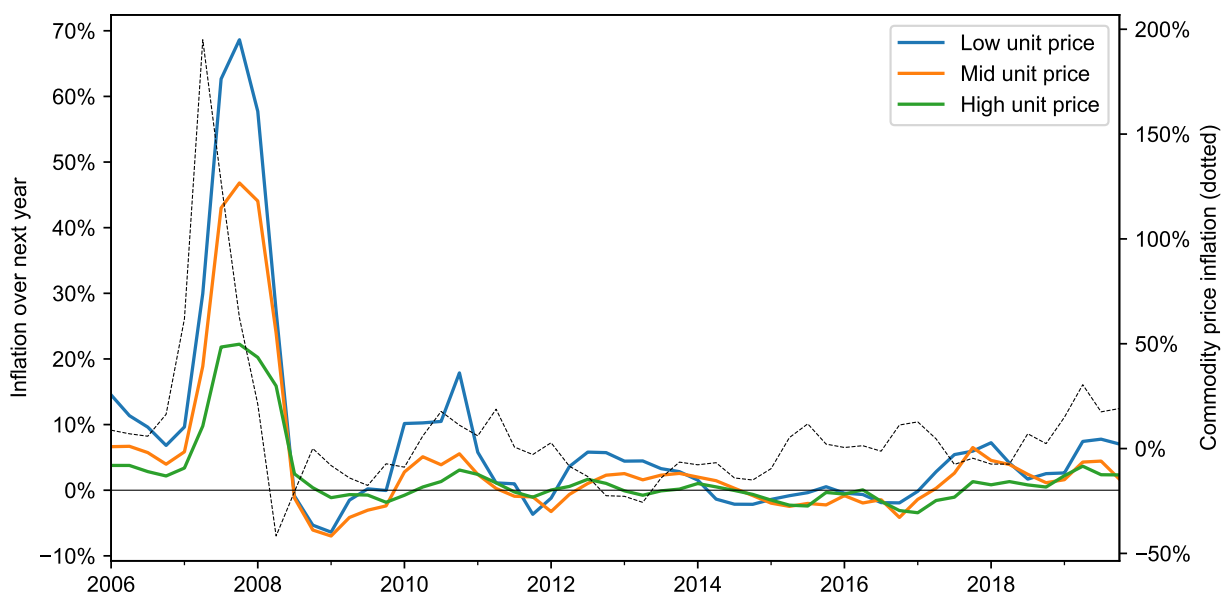
(a) Inflation (in percentages)
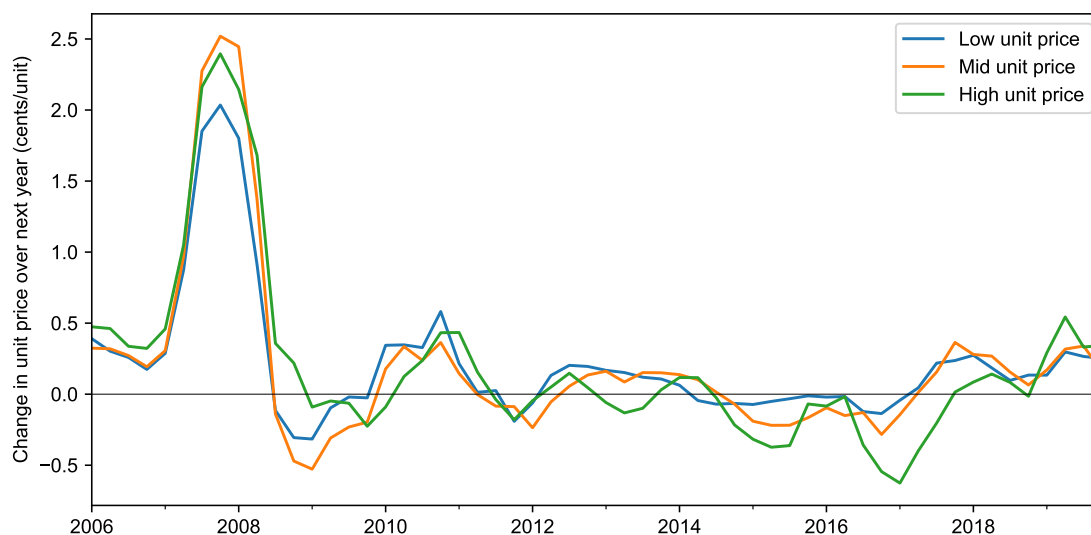
(b) Price change (in levels)

Note: Both panels plot price changes for rice products in the Nielsen scanner data. In each quarter, all UPCs are separated into three groups with equal quarterly sales by (quantity-weighted) average unit price over the prior year. Panel (a) plots the sales-weighted average inflation rate over the next year for products in each group, alongside commodity rice inflation. Panel (b) plots the sales-weighted average change in price levels over the next year for products in each group.

these three groups of rice products. As shown in the top panel, a run-up in rice commodity prices into 2008 led to much higher inflation for rice products with lower unit prices—the average inflation rate for low unit price rice products reached nearly 70 percent in 2008, compared to under 25 percent for high unit price products.[30] These differences disappear when comparing the price changes in levels across unit price groups in the bottom panel: products in all unit price groups had roughly the same increase in absolute prices.

To test how pass-through in logs and levels varies in the cross-section of products, Table 6 reports estimates from the following specifications,

$$\Delta \pi_{i,t} = \alpha_i + \sum_{g=1}^{3} \beta_g \left(1\{G(i,t) = g\} \times \pi_t^c \right) + \varepsilon_{i,t}, \tag{5}$$

$$\Delta p_{i,t} = \alpha_i + \sum_{g=1}^{3} \beta_g \left(1\{G(i,t) = g\} \times \Delta c_t \right) + \varepsilon_{i,t}, \tag{6}$$

where $G(i,t) \in \{1,2,3\}$ is the unit price group of product i in quarter t, and $\pi_t^c = (c_{t+4}/c_t) - 1$ and $\Delta c_t = c_{t+4} - c_t$ are changes in commodity prices over the next year in logs and levels.[31]

Across product groups, panel A shows that retail price inflation is most sensitive to commodity price inflation for products in the lowest unit price group. The sensitivity of retail price inflation to commodity inflation systematically declines with unit price across all three product categories (rice, flour, and coffee). In contrast, Panel B shows that there are no systematic differences in the sensitivity of retail price changes to commodity price changes *in levels* across unit price groups, consistent with complete pass-through in levels across products explaining heterogeneity in log pass-through.[32]

Exploiting variation in identical products' prices across retailers. Heterogeneity in unit prices across products in a category can come from differences in markups or from differences in non-commodity input costs. Even when firms use multiplicative markup rules, log pass-through can decline with unit price if the heterogeneity in unit prices across

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[30]Childs and Kiawu (2009) provide a detailed account of the factors leading to the rise in rice prices in 2008. The run-up was prompted by adverse weather shocks to wheat-growing areas from 2006–2008, and subsequent trade restrictions by Vietnam, India, and other major rice-exporting countries to ensure adequate rice supply for their domestic markets.

[31]Specifications (5) and (6) do not measure long-run pass-through. Estimating long-run pass-through in logs and levels across the unit price groups of products yields similar qualitative patterns—log pass-through decreases with unit price, while pass-through in levels is about constant with unit price—but the results are noisier since long-run pass-through cumulates over several lags.

[32]Appendix Table A5 shows similar results if we instead split products into five unit price groups.

Table 6: Higher-priced products exhibit lower log pass-through, with no systematic difference in level pass-through.

Panel A: In percentages

	Retail price inflation		
	Rice	Flour	Coffee
Commodity Inflation	0.226**	0.074**	0.110**
	(0.019)	(0.005)	(0.014)
Commodity Inflation × Unit Price Group 2	−0.075**	−0.007	−0.064**
	(0.014)	(0.009)	(0.015)
Commodity Inflation × Unit Price Group 3	−0.150**	−0.046**	−0.091**
	(0.022)	(0.009)	(0.017)
UPC FEs	Yes	Yes	Yes
N (thousands)	399.4	101.4	1570.0
R^2	0.15	0.05	0.14

Panel B: In levels

	Δ Retail price		
	Rice	Flour	Coffee
Δ Commodity Price	0.057**	0.035**	0.059**
	(0.007)	(0.007)	(0.008)
Δ Commodity Price × Unit Price Group 2	0.007	0.009	−0.020
	(0.006)	(0.013)	(0.013)
Δ Commodity Price × Unit Price Group 3	0.005	−0.023	−0.029*
	(0.012)	(0.015)	(0.017)
UPC FEs	Yes	Yes	Yes
N (thousands)	399.4	101.4	1570.0
R^2	0.07	0.05	0.16

Note: Panel A reports results from specification (5), and panel B reports results from specification (6). The three columns use products from rice, all purpose white flour, and roasted coffee, respectively. In each quarter, each retailer-UPC pair is assigned to three groups with equal sales by (quantity-weighted) average unit price over the past year. Unit Price Groups 2–3 are indicators for whether a retailer-UPC pair is assigned to the mid- or high-unit price group. Standard errors clustered by brand. * indicates significance at 10%, ** at 5%.

Figure 7: Price of same coffee UPC in two stores in same 3-digit ZIP in Philadelphia, PA.
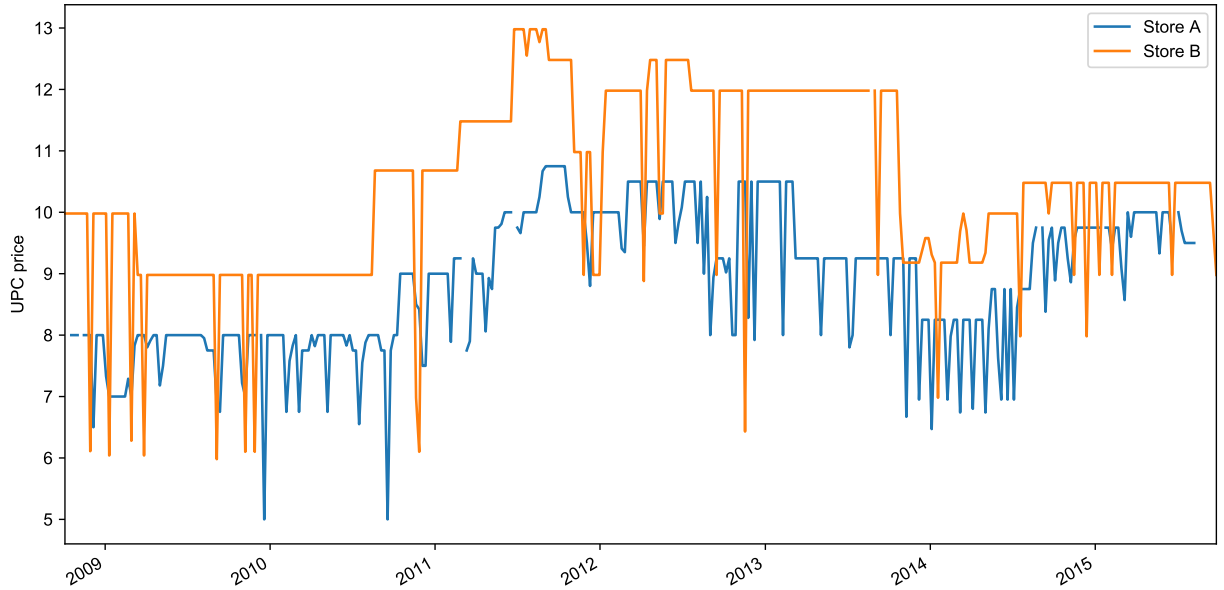
products largely steps from differences in non-commodity input costs.[33]

To narrow in on how heterogeneity in markups affects pass-through, I exploit the fact that different retailers often sell the same product at different prices. If differences in prices charged for the same product primarily reflect different retail markups, rather than differences in costs across retailers, differences in pass-through for the same product across retailers will isolate the effect of markups on pass-through.

To fix ideas, consider two retailers selling the same UPC, one with a low markup (store A) and one with a high markup (store B). Figure 7 shows, for example, the price of a single coffee UPC at two different stores in the same three-digit ZIP code in Philadelphia, PA. Excluding some temporary sales, store A consistently charges a lower price than store B.

If both stores A and B choose a fixed percentage rule, when the cost of the UPC rises, store B (the retailer with the higher markup) should increase its price by more in levels. On the other hand, if both stores choose a fixed, additive unit margin, when the cost of UPC rises, the absolute price change in both store A and store B should be similar, and the price change in percentage terms for store B should be lower.

Denote the quantity-weighted average price of UPC i sold by retailer r in quarter t by $p_{i,r,t}$. We can test these predictions using the specification,

$$\Delta p_{i,r,t} = \gamma \overline{\Delta p_{i,t}} + \beta \left(\mu_{i,r,t} \times \overline{\Delta p_{i,t}} \right) + \delta \mu_{i,r,t} + \alpha_i + \varepsilon_{i,r,t}. \tag{7}$$

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[33]Of course, if firms use multiplicative markup rules, pass-through in levels should still increase with unit price if unit prices are positively correlated with markups, which is rejected in Table 6.

where $\Delta p_{i,r,t} = p_{i,r,t+4} - p_{i,r,t}$ is the year-over-year change in the price charged by retailer r for UPC i starting in quarter t, $\overline{\Delta p_{i,t}}$ is the average year-over-year change in the price charged by all retailers for UPC i starting in quarter t, $\mu_{i,r,t}$ is a measure of the markup charged by retailer r for UPC i, and α_i are UPC fixed effects. If retailers choose constant percentage markups, then high-markup retailers should increase their prices more than other retailers when the cost of UPC i increases, and therefore we should find $\beta > 0$. On the other hand, if retailers set constant additive unit margins, we should find $\beta \approx 0$.

Since we cannot observe retail markups directly, I use the log deviation in the price retailer r charges for UPC i relative to the average price of UPC i as an empirical proxy for $\mu_{i,r,t}$:

$$\widehat{\mu}_{i,r,t} = \log\left(p_{i,r,t}/\overline{p_{i,t}}\right).$$

Panel A of Table 7 reports results from (7) for rice, flour, and coffee products. For all three categories, the estimated coefficient β is slightly negative and indistinguishable from zero at the 5 percent level (columns 1–3). That is, retailers selling the same UPC with different markups exhibit similar price changes in levels. These results are robust to estimating β using UPC-time fixed effects (columns 4–6).

Panel B of Table 7 reports results from the analogous specification measuring price changes in percentage terms,

$$\pi_{i,r,t} = \tilde{\gamma}\overline{\pi_{i,t}} + \tilde{\beta}\left(\mu_{i,r,t} \times \overline{\pi_{i,t}}\right) + \tilde{\delta}\mu_{i,r,t} + \tilde{\alpha}_i + \varepsilon_{i,r,t}, \tag{8}$$

where $\pi_{i,r,t} = (p_{i,r,t+4}/p_{i,r,t}) - 1$ is the year-over-year percentage change in the price charged by retailer r for UPC i starting in quarter t, and $\overline{\pi_{i,t}}$ is the average year-over-year percent change in the price charged by all retailers for UPC i starting in quarter t. In all cases, the estimated coefficient $\tilde{\beta} < 0$: retailers with higher markups increase prices by less in percentage terms than retailers with lower markups.[34]

Thus, by exploiting variation in markups across retailers, we find that products have similar pass-through in levels, which appears as lower log pass-through for high markup products. These findings are consistent with the additive margin pricing model, rather than the multiplicative markup pricing model.

Since this approach does not require information on upstream commodity costs, we can extend this analysis to a broader set of product modules in the data. I estimate

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[34]The fact that the estimated interaction coefficient, $\tilde{\beta}$, is close to -1 in all specifications in Table 7 is predicted by the fixed unit margin model. Suppose retailer r sets price $p_{i,r} = c_i + \alpha_{i,r}$. Following a change in the product cost c_i, the retailer's price change in percentage terms is approximately $d\log p_{i,r} \approx c_i/(c_i + \alpha_{i,r})d\log c_i$. If all other retailers also charge fixed additive unit margins, the change in the average price $\bar{p}_i$ is approximately, $d\log\bar{p}_i \approx (c_i/\bar{p}_i)d\log c_i$. Combining yields $d\log p_{i,r} \approx (1 - \log(p_{i,r}/\bar{p}_i))d\log\bar{p}_i$.

Table 7: Exploiting variation in markups for identical products across retailers.

Panel A: In levels

			ΔUPC Price at Retailer			
	Rice	Flour	Coffee	Rice	Flour	Coffee
	(1)	(2)	(3)	(4)	(5)	(6)
Avg. ΔUPC Price	0.815**	0.853**	0.909**			
	(0.048)	(0.064)	(0.018)			
Avg. ΔUPC Price $\times \log(p_{i,r,t}/\overline{p_{i,t}})$	−0.243*	−0.321	−0.080	−0.019	−0.200	−0.123
	(0.140)	(0.286)	(0.192)	(0.111)	(0.216)	(0.352)
UPC FEs	Yes	Yes	Yes			
UPC-Time FEs				Yes	Yes	Yes
N (thousands)	399.4	101.4	1570.0	399.4	101.4	1570.0
R^2	0.38	0.41	0.49	0.51	0.50	0.55

Panel B: In percentages

			UPC Inflation at Retailer			
	Rice	Flour	Coffee	Rice	Flour	Coffee
	(1)	(2)	(3)	(4)	(5)	(6)
UPC Avg. Inflation	0.895**	0.943**	0.915**			
	(0.042)	(0.035)	(0.013)			
UPC Avg. Infl. $\times \log(p_{i,r,t}/\overline{p_{i,t}})$	−0.841**	−0.853**	−0.958**	−0.988**	−0.879**	−1.386**
	(0.052)	(0.177)	(0.084)	(0.104)	(0.250)	(0.213)
UPC FEs	Yes	Yes	Yes			
UPC-Time FEs				Yes	Yes	Yes
N (thousands)	399.4	101.4	1570.0	399.4	101.4	1570.0
R^2	0.58	0.55	0.53	0.64	0.60	0.58

Note: Panel A reports results from specification (7), and panel B reports results from specification (8). The latter three columns add UPC-quarter fixed effects. The average change in the UPC price and the average UPC inflation are sales-weighted averages over all stores in the Nielsen Retail Scanner dataset, which includes some stores that cannot be assigned to a retail chain. $\log(p_{i,r,t}/\overline{p_{i,t}})$ denotes the log deviation of the (quantity-weighted) average price for UPC i at retail chain r in quarter t from UPC i's average price across all stores in quarter t. Standard errors clustered by brand. * indicates significance at 10%, ** at 5%.

Table 8: Exploiting variation in markups for identical products across retailers: Summary of results across all product modules.

Share of modules	Unweighted	Observations-weighted	Sales-weighted
Panel A: In levels			
Positive coefficient	9.7	5.6	7.2
Not significant	63.2	54.1	54.0
Negative coefficient	27.1	40.3	38.8
Panel B: In logs			
Positive coefficient	0.6	0.0	0.0
Not significant	13.7	3.6	3.6
Negative coefficient	85.7	96.4	96.4

Note: Summary of results from specifications (7) (for panel A) and (8) (for panel B) estimated across 616 product modules. Each cell reports the fraction of product modules for which the estimated interaction between the average UPC price change (in levels or logs) and the relative price of the product at the retailer is significant at a 5% level. Driscoll-Kraay standard errors used in all specifications.

specifications (7) and (8) for each product module and report the share of modules with significant positive and negative interaction coefficients in Table 8.[35] Fewer than 10 percent of product modules exhibit the positive interaction between retailer markups and inflation sensitivity that would be predicted by constant-markup pricing. For over 85 percent of product modules (accounting for over 96 percent of sales), the same product has significantly lower log pass-through at retailers where it is sold at a higher markup.[36]

Other explanations of heterogeneous log pass-through. Previous work proposes other explanations for the heterogeneity in log pass-through across products. Two of the most popular attribute variation in log pass-through to heterogeneity in market share or firm size—firms with larger market shares exhibit lower log pass-through in models of nested oligopoly (such as Atkeson and Burstein 2008) or when firms' residual demand curves are log-concave (see Melitz 2018, Matsuyama and Ushchev 2022)—or to differences in the price sensitivity of firms' customers over the business cycle (e.g. Li 2019).

Appendix Tables A6-A8 suggest that complete pass-through in levels has greater explanatory power for cross-sectional heterogeneity in pass-through than these two alterna-

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[35]This analysis uses the 616 product modules in Nielsen departments that fall within the food-at-home basket (departments for dry grocery, frozen foods, dairy, deli, packaged meat, and produce) and that have at least 250 observations (the median product module has about 100,000 observations).

[36]As a robustness exercise, Appendix Table A9 reports results analogous to Table 8, but instead using a leave-one-out measure of the average UPC price change and UPC inflation across retailers. Under that specification, in over 50 percent of product modules (accounting for nearly 80 percent of total sales), products at high-margin retailers exhibit significantly lower log pass-through.

tives. After controlling for unit prices, products' market shares and the average income of a products' buyer are insignificant predictors of variation in pass-through across rice, flour, and coffee products. While not a comprehensive test of all other mechanisms proposed by the literature, these results suggests that complete pass-through in levels and variation in margins provides a better fit for the cross-sectional variation in log pass-through observed in the data than other popular models.

5 Explaining Complete Pass-Through in Levels

Why do firms in the studied industries exhibit complete pass-through in levels? Two candidate explanations—perfect competition and non-isoelastic demand—could in principle explain complete pass-through in levels, but generate predictions that are at odds with other features of the data.

I propose a different micro-foundation for complete pass-through in levels. In particular, a combination of overhead costs with a *safety margin constraint*—an aversion to variable profits falling below overhead costs—can generate pass-though in levels. This model also yields predictions for the behavior of entry and gross margins that differ from the workhorse Dixit and Stiglitz (1977) model, which I test at the close of this section.

5.1 Is Complete Pass-Through in Levels Due to Perfect Competition?

One explanation for complete pass-through in levels is that firms in the studied industries set prices equal to marginal cost. This implies that the markup $\mu = 1$, and hence pass-through is complete in levels.

Perfect competition is at odds with several other features of the data, however. In the case of retail gasoline, perfect competition is difficult to square with the presence of price cycles (shown in Figure 4) and the presence of substantial price dispersion even within postcodes or narrowly defined neighborhoods. Moreover, previous work by Wang (2009b)—who collects sales data from seven gas stations in Perth—estimates elasticities of demand facing firms between 6 and 19. While high, these price elasticities imply firm-level demand curves that are far from horizontal.

Table 9 shows evidence of substantial markups over marginal cost in the studied food products as well. First, I calculate retail markups as average retail prices over wholesale costs for matched UPCs.[37] Second, I use Hausman (1996) instruments to estimate elas-

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[37]These wholesale costs come from PromoData Price-Trak. Sangani (2022) provides a detailed overview of these wholesale cost data, and I report median retail markups constructed in a manner identical to the

Table 9: Evidence on markups in studied food products.

Markup (Median [Q1 Q3])	Coffee	Rice	Flour
1. Retail markup	14% [3%, 35%]	49% [33%, 72%]	24% [14%, 34%]
2. Demand elasticity	40% [26%, 72%]	90% [30%, –]	206% [57%, –]
3. Literature	58%[a]	43–75%[b]	n/a

Note: Retail markups for each UPC are calculated as the quantity-weighted average price over PromoData wholesale cost. Demand elasticity markup estimates are constructed using the standard Lerner index, with demand elasticities estimated using a Hausman (1996) instrument described in Appendix D. Literature estimates are from ([a]) Nakamura and Zerom (2010) and ([b]) Park (2013).

ticities of demand for each UPC at each store, and calculate implied markups using the usual Lerner formula. (The details on estimating these demand elasticities are relegated to Appendix D.) Finally, I draw from previous literature that has estimated markups in these markets (Nakamura and Zerom 2010, Park 2013). All three measures suggest substantial markups over marginal cost, inconsistent with perfect competition.

5.2   Is Complete Pass-Through in Levels Due to Demand Curvature?

Pass-through in levels may be complete if the demand curves facing firms are log-concave, rather than isoelastic, resulting in percent markup adjustments that coincide with a fixed additive margin. Appendix B shows that pass-through is complete in levels if the super-elasticity of demand is exactly equal to one (i.e., $\frac{\partial \log \sigma}{\partial \log p} = 1$, where $\sigma = -\frac{\partial \log D(p)}{\partial \log p}$). As shown by Bulow and Pfleiderer (1983), semilog demand curves (i.e., demand curves of the form $D(p) = a - b \log p$) satisfy this requirement.

Burya and Mishra (2023) develop a technique to measure the super-elasticity of demand directly in the data. In particular, they show that by estimating the specification,

$$\log q_{i,t} = \eta \log p_{i,t} + \kappa (\log p_{i,t})^2 + \gamma X_{i,t} + \varepsilon_{i,t},$$

the ratio κ/η provides an estimate of the super-elasticity of demand. I adopt their approach, using a Hausman (1996) instrument for prices to identify changes in prices that are plausibly exogenous to local demand conditions, to estimate the super-elasticity of demand individually for each UPC in each store in the coffee, rice, and flour product

retail markups studied in Sangani (2022).

Table 10: Share of store-UPC pairs with estimated super-elasticity of demand below one.

Percent of store-UPC pairs	Coffee	Rice	Flour
Point estimate below one	90.0	98.7	79.2
Reject over one at $p = 0.05$	62.6	89.2	52.5
N (thousands)	18.6	9.7	7.1

categories. This estimation procedure is described in detail in Appendix D.

Table 10 reports the share of store-UPC pairs in each product category where the estimated super-elasticity of demand is below one. In the vast majority of cases, estimated super-elasticities of demand fall short of one, and super-elasticities of at least one are rejected at a five percent level in over half of the store-UPC observations. These estimates suggest that demand curves in the data are not sufficiently concave to generate pass-through in levels.

Relaxing Leontief, constant returns, and uncorrelated costs. What if we relax the other restrictions we have imposed so far, such as constant returns and Leontief production? In Appendix B, I allow production to take the more general form,

$$y = \left(\omega x^{\frac{\theta-1}{\theta}} + (1-\omega)\ell^{\alpha\frac{\theta-1}{\theta}} \right)^{\frac{\theta}{\theta-1}},$$

where θ is the elasticity of substitution across inputs, ω are weights in production on the two inputs, and $\alpha < 1$ generates decreasing returns to scale in the non-commodity input. I show that relaxing any one assumption—i.e., relaxing Leontief production, constant returns to scale, or uncorrelated other variable costs—requires knife-edge conditions to deliver complete pass-through in levels that are unlikely to hold in practice.

5.3   Price and Entry Dynamics in a Model with Safety Margins

The model consists of a continuum of firms that produce varieties of an output good using two variable inputs: a commodity x and labor ℓ. Production is Leontief and constant returns to scale, so that each firm's output is given by

$$y = \min\{x, \ell\}.$$

Denote the per-unit commodity cost by c and the wage (the per-unit labor cost) by w.

In addition to variable commodity and labor costs, firms have an overhead cost each period, which is paid in units of labor. I assume each firm's overhead cost is given by $wf_oN^{-\zeta}$. Here, N is the mass of firms in the market, and $-\zeta$ is the elasticity of the overhead cost to the mass of firms. I assume that $\zeta > 0$, so that overhead costs fall with the mass of firms in the market. This assumption is made to reflect the fact that overhead costs may scale with expected capacity (for instance, a small number of retailers serving a large market may have to pay higher rent costs for larger floor space). As we will see below, $\zeta > 0$ is not necessary in the standard multiplicative-markup model, but is necessary to close the equilibrium with free entry in the additive margin model.

Thus, the total cost function for each firm is given by

$$C(y) = (c + w)y + wf_oN^{-\zeta},$$

and profits are given by

$$\pi(p) = pD(p) - C(D(p)).$$

Firms face downward-sloping, isoelastic residual demand curves. The demand curve for each retailer depends on its price p, the aggregate market price P, the elasticity of substitution across retailers' outputs $\sigma > 1$, the mass of competing firms N, and a demand shock ε,

$$D(p) = \varepsilon \frac{1}{N} \left(\frac{p}{P}\right)^{-\sigma}. \tag{9}$$

I assume that ε has an expected value of one and is drawn from a distribution with CDF H. Crucially, the demand shock ε introduces a source of uncertainty for the firm. As we will see below, each firm sets its price before the shock ε is realized. Hence, while the firm is correct about its profits in expectation, the level of demand may fluctuate around its expected value.[38]

In the face of this uncertainty, I introduce a *safety margin constraint* by which firms bound the risk that their profits fall below some pre-specified level. This safety margin constraint was first discussed by Fellner (1948) and was later formalized by Day et al. (1971). In particular, this form of the safety margin constraint assumes that retailers maximize expected profits subject to the constraint that the probability of profits fall

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[38]Total industry demand in (9) is constant and normalized to one. Aggregate demand for retail gasoline and staple food products tends to be inelastic (though elasticities of substitution between firms within the industry may be large). For example, research by the USDA estimates the elasticities of aggregate demand for flour, rice, and coffee to be 0.07, −0.07, and −0.12, respectively (Okrent and Alston 2012).

below zero is at most ϕ,[39]

$$\Pr\left[\pi(p) \le 0\right] \le \phi. \tag{10}$$

Thus, the firm's problem is to choose an output price to maximize expected profits, subject to the residual demand curve (9) and to the safety margin constraint (10).

Finally, the model is closed with a standard zero profit condition. Firms pay a cost wf_e to enter in each period, and entry occurs until the sum of discounted expected future profits is equal to the entry cost:

$$\mathbb{E}_t \sum_{k=0}^{\infty} \beta^k \pi_{t+k} = w_t f_e, \qquad \text{for all } t. \tag{11}$$

I assume that wages are expected to grow at a constant rate g (i.e., $\mathbb{E}_t[w_{t+k}] = (1+g)^k w_t$), and that the ratio of the commodity cost to the wage, c_t/w_t, is a random walk with zero drift. For ease of notation, define $\Delta \equiv (1 - \beta(1+g))$.

Definition 1 (Symmetric equilibrium). Given a path for commodity prices and wages (c_t, w_t), a symmetric equilibrium is a set (p_t, P_t, N_t) such that the price chosen by firms in each period, p_t, maximizes expected profits subject to (9) and (10), the aggregate market price is equal to the price set by all firms, $P_t = p_t$, and the free entry condition (11) is satisfied in all periods.

We now move to characterizing prices and the mass of firms in equilibrium. First, we consider periods where the safety margin constraint (10) does not bind. Proposition 1 shows that the price is given by a multiplicative markup over marginal cost, as in the workhorse Dixit and Stiglitz (1977) model.

Proposition 1 (Dixit and Stiglitz 1977). *In periods where the safety constraint does not bind, the output price is given by*

$$p_t^{DS} = \frac{\sigma}{\sigma - 1}(c_t + w_t).$$

Across these periods, the mass of firms is positively correlated with the commodity cost, $dN_t^{DS}/(d(c_t/w_t)) > 0$, and gross margins are constant at $m_t^{DS} = 1/\sigma$.

Proposition 1 describes properties of the well-known Dixit and Stiglitz (1977) equilibrium: prices are set equal to a fixed, multiplicative markup over marginal cost and gross margins are constant. Since markups are fixed in percentage terms, firms' per-unit

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[39]This constraint is referred to as "strict safety-first" pricing by Day et al. (1971). This assumption has also been recently used by Altomonte et al. (2015), who consider this safety margin constraint as a rationale for full-cost pricing. As I show below, while this safety margin constraint implies that firms factor overhead costs into their pricing decisions, it does not necessarily imply that firms will price at average cost.

variable profits increase when commodity costs rise. Thus, new firms enter until variable profits fall to meet the zero profit condition.

Next, Proposition 2 describes the dynamics of prices, entry, and gross margins when the safety constraint binds.

Proposition 2 (Safety margin pricing). *In periods where the safety constraint binds, the mass of firms $N_t^{safe} = N^{safe}$ is constant, and the output price is given by*

$$p_t^{safe} = c_t + w_t + \alpha w_t f_o,$$

where the constant $\alpha > 1$. Across these periods, gross margins are negatively correlated with the commodity cost, $dm_t^{safe}/d(c_t/w_t) < 0$. Holding all parameters fixed, $p_t^{safe} \geq p_t^{DS}$ and $N_t^{safe} \geq N_t^{DS}$.

Rather than setting a multiplicative markup, when the safety constraint binds firms instead set price equal to a fixed, additive margin over marginal. This margin is related to the overhead cost, since it is chosen to ensure that firms bound the risk of being unable to meet their overhead costs.

In contrast to the Dixit and Stiglitz (1977) model, in periods where the safety constraint binds, the mass of firms is unrelated to the commodity cost. Intuitively, since firms' variable profits are now unrelated to the commodity cost, changes in the commodity cost do not incentivize new entry or exit. Gross margins, on the other hand, fall when the commodity cost rises, since complete pass-through in levels results in incomplete log pass-through.

When will price and entry dynamics resemble Proposition 1 versus Proposition 2? Proposition 3 shows that which pricing regime an industry falls into depends on the level of commodity costs relative to wages, as well as parameters such as the overhead cost and the cost share of commodity inputs.

Proposition 3 (When constraint binds). *There exists a cutoff c^* such that the safety margin constraint binds whenever $c_t/w_t \leq c^*$ and does not bind otherwise. The cutoff c^* is increasing in f_o, decreasing in Δf_e when $\zeta \in [0,1)$, decreasing in ϕ, and increasing in the variance of ε.*

The comparative statics in Proposition 3 are informative about the types of industries where price and entry dynamics are more likely to follow Proposition 2. In particular, industries are more likely to resemble Proposition 2 when overhead costs are high and demand is volatile. Intuitively, these are the industries where fluctuations in commodity costs increase the risk that overhead costs are not covered by variable profits. The safety margin constraint is also more likely to bind if the risk of negative profits that retailers are willing to tolerate is low or if entry costs are low.

Figure 8: Wheat commodity costs and downstream (bread industry) gross margins.
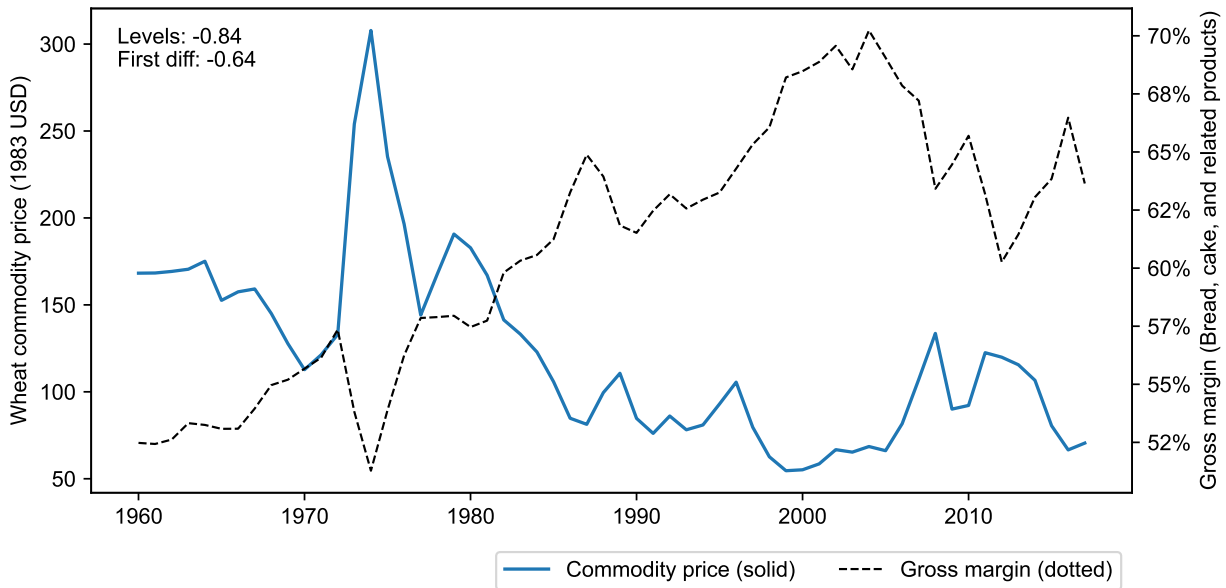
Note: Gross margins are total sales minus costs of goods sold (material costs) as a share of sales, from the NBER-CES manufacturing database. Wheat commodity prices are from UNCTADSTAT, deflated to 1983 dollars using CPI excluding food and energy.

5.4 Empirical Evidence on Entry and Margins

I test how the behavior of entry and gross margins in the data compare to the predictions from Propositions 1 and 2. First, I test how the mass of firms changes with commodity costs, using the retail gasoline and food products markets studied earlier. Appendix Figure A12 plots the number of retail gas stations and the wholesale gas price from 2003–2023 in Perth. There is little responsiveness of entry to commodity costs: the correlation of month-over-month changes in the number of firms with changes in the wholesale gas price is 0.07 (statistically indistinguishable from zero). For food products, it is difficult to provide an exact count of the number of firms in the market. As a proxy, I plot the market shares of the largest brands in the coffee, flour, and rice product categories in Appendix Figure A13. If the number of firms were positively correlated with commodity costs, the market share of top brands should erode when commodity costs increase. Instead, the data show no discernible relationship between commodity costs and market shares of top brands. Hence, the data on entry dynamics are more consistent with the predictions of the safety margin regime (Proposition 2) than the workhorse model.

Second, I test how industry gross margins correlate with upstream commodity costs. For this analysis, I use data on gross margins for manufacturing industries from the NBER-CES manufacturing industry database (Becker et al. 2021). These data include sales, cost

Table 11: Correlation between commodity costs and downstream gross margins.

	Gross margins from SIC industry		Correlation Costs = Materials		Correlation Mat., Energy, Labor	
Commodity	Description	SIC	Levels	First diff.	Levels	First diff.
Sugar	Candy and confectionery products	2064	-0.58**	-0.37**	-0.51**	-0.22*
Beef	Sausages and other prepared meats	2013	-0.82**	-0.39**	-0.82**	-0.30**
Wheat	Flour and other grain mill products	2041	-0.80**	-0.55**	-0.73**	-0.46**
Wheat	Prepared flour mixes and doughs	2045	-0.80**	-0.57**	-0.79**	-0.50**
Wheat	Bread, cake, and related products	2051	-0.84**	-0.64**	-0.80**	-0.54**
Rice	Rice milling	2044	-0.70**	-0.17	-0.60**	-0.09
Coffee	Roasted coffee	2095	-0.79**	-0.58**	-0.78**	-0.56**
Cocoa beans	Chocolate and cocoa products	2066	-0.36**	-0.07	-0.35**	-0.03
Cotton	Broadwoven fabric mills, cotton	2211	0.02	-0.42**	-0.66**	-0.39**
Milk	Cheese; natural and processed	2022	-0.66**	-0.61**	-0.56**	-0.50**
Milk	Dry, condensed, evaporated products	2023	-0.52**	-0.58**	-0.51**	-0.53**
Aluminum	Aluminum sheet, plate, and foil	3353	-0.73**	-0.41**	-0.71**	-0.30*
Aluminum	Aluminum die-castings	3363	-0.63**	-0.57**	-0.63**	-0.25
Orange juice	Frozen fruits and vegetables	2037	-0.63**	-0.18	-0.70**	-0.17

Note: Industry data from NBER-CES manufacturing database (1958–2018). Variable costs defined as material costs or sum of material, energy, and production labor costs. Commodity prices are from UNCTADSTAT (1960–2017), except milk, aluminum, and frozen orange juice, which are from the IMF Commodities database (1980–2018). Commodity prices deflated using core CPI. * indicates significance at 10%, ** at 5%.

of goods sold (materials costs), labor costs, and other aggregated industry-level data for four-digit SIC industries over 1958–2018. Thirteen of these industries can be clearly paired with an upstream commodity; for example, I pair candy and other confectionary products (SIC 2064) with sugar commodity prices, and so on.[40]

As an example, Figure 8 plots gross margins (measured as sales minus costs of goods sold, as a percent of sales) for SIC industry 2051 (bread, cake, and related products) alongside the wheat commodity price from 1960–2017. Consistent with Proposition 2, there is a negative correlation both in levels ($\rho = -0.84$) and in first differences ($\rho = -0.64$) for the two series. Table 11 extends this analysis to all thirteen SIC industries that can be paired with an upstream commodity. In nearly all cases, we find a negative correlation between upstream commodity prices and downstream industry gross margins, whether the measure of variable costs includes or excludes labor and whether we calculate the correlation in levels or in first differences.

Both the behavior of entry and gross margins are consistent with the predictions of

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[40]To take advantage of the long time span over which these data are available, I use annual commodity prices from UNCTADSTAT where possible to extend the analysis back to 1960. UNCTADSTAT does not contain commodity price data for milk, aluminum, and frozen orange juice prices, so for industries using those commodities I instead use the IMF commodity price data that were used in Section 4.

safety margin pricing (Proposition 2) rather than the workhorse Dixit and Stiglitz (1977) model (Proposition 1).

6   The Unequal Incidence of Commodity Shocks

This section explores the implications of complete pass-through in levels for inflation across the income distribution. As we've seen, increases in upstream costs result in greater inflation for low-markup products within a category. Since low-income households disproportionately purchase low-price and low-markup products (Sangani 2022), inflation rates for low-income households will be more sensitive to upstream commodity prices even within narrow product categories.

Section 6.1 shows that inflation rates for low-income households are indeed more sensitive to commodity costs in granular product categories like coffee, rice, and flour. Section 6.2 then extends the analysis to the entire food-at-home bundle.

6.1   Within-Category Inflation Inequality

Section 4 showed that log pass-through is higher for low unit-price products within a category. To the degree that differences in unit prices are driven by differences in markups, we should expect log pass-through to fall with the level of markups and—if high-income consumers tend to purchase higher markup products—to fall with buyer income.

These predictions are borne out in Table 12, which explores how the sensitivity of retail price inflation for products in a category to commodity inflation varies with products' retail markups and buyer income. Retail markups are measured using wholesale cost data from PromoData Price-Trak following Sangani (2022).[41] I measure the average buyer income for each UPC in each quarter as the sales-weighted average of buyers' incomes across all observed transactions of the UPC in the Nielsen Homescan panel. Across rice, flour, and coffee products, the sensitivity of retail price inflation to commodity inflation falls with products' retail markups and thus with average buyer income.

I now turn to quantifying differences in category-level inflation rates experienced by households over the income distribution. I calculate the inflation rate experienced by households in quintile j, π_t^j, as the expenditure-weighted average of inflation rates on

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[41]See Appendix A of Sangani (2022) for a detailed overview. Since product coverage by PromoData varies substantially from year to year, I use data from a single year (2008) for which the match rate of UPCs from PromoData to the Nielsen scanner data is highest to calculate an average retail markup for each UPC.

Table 12: Products with high retail markups and high-income buyers have lower log pass-through.

	Rice		Flour		Coffee	
Retail product inflation	(1)	(2)	(3)	(4)	(5)	(6)
Commodity Inflation	0.155**	0.149**	0.053**	0.054**	0.104**	0.057**
	(0.018)	(0.016)	(0.011)	(0.008)	(0.011)	(0.012)
Comm. Infl. × Log(Retail Markup)	−0.207**		−0.119**		−0.198**	
	(0.078)		(0.049)		(0.048)	
Comm. Infl. × Log(Buyer Income)		−0.135**		−0.114**		−0.219**
		(0.043)		(0.036)		(0.049)
UPC FEs	Yes	Yes	Yes	Yes	Yes	Yes
N (thousands)	197.7	329.8	64.7	79.6	253.3	1269.3
R^2	0.11	0.13	0.02	0.03	0.11	0.13

Note: Standard errors clustered by brand. * indicates significance at 10%, ** at 5%.

each individual product i purchased by households in quintile j,

$$\pi_t^j = \frac{\sum_i \lambda_{i,t}^j \pi_{i,t}}{\sum_i \lambda_{i,t}^j},$$

where $\lambda_{i,t}^j$ are the total expenditures on product i by households in quintile j at time t and $\pi_{i,t}$ is the inflation rate of product i over the next year starting in quarter t. Note that these measures of inflation use expenditure shares at time t, and hence are not contaminated by how households substitute across products in response to price changes.

As an example, Figure 1 (in the Introduction) plots the difference between the inflation rate for coffee products for households in the lowest income quintile and the highest income quintile. There are large swings in the extent of the within-category inflation rates, with spikes in 2011 and 2014 coinciding with increases in coffee commodity costs. The inflation gap is positive on average, consistent with the secular drivers of inflation inequality documented by Jaravel (2019, 2021), but features cyclical swings well described by complete pass-through in levels. Note that this cyclical source of inflation inequality means that during periods of commodity price deflation, such as in 2012–2013, the inflation gap between low-income and high-income households can even be negative.

Figure 9 shows the log pass-through of commodity costs for each income quintile in each product category (flour, rice, and coffee). Intuitively, differences in log pass-

Figure 9: Log pass-through of commodity costs by income quintile.

(a) Flour. **(b)** Rice. **(c)** Coffee.
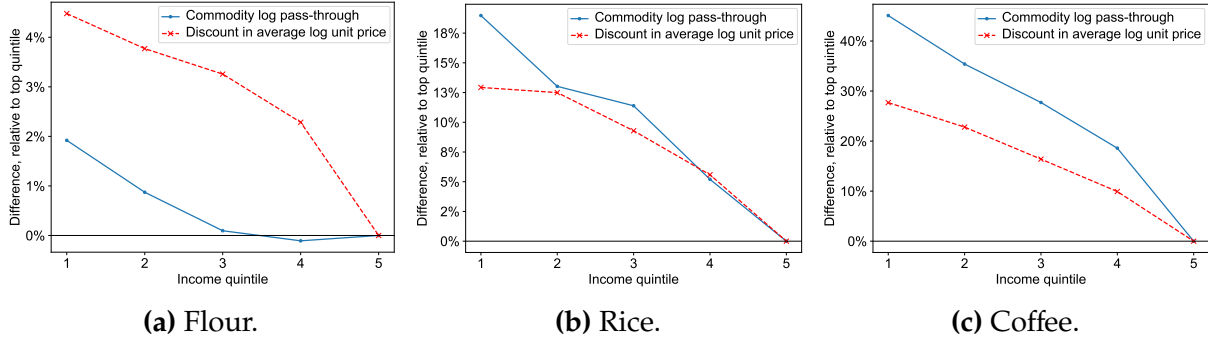
Note: Discount in average log unit price shows the percent difference in the average, posted unit price of products consumed by households in the income group relative to the highest income group (e.g., households in the lowest income quintile buy flour products that have a 4.3% lower unit price on average than flour products bought by households in the highest income quintile). Commodity log pass-through shows the percentage difference in the long-run log pass-through of commodity costs to retail prices for households in the income group vs. the highest income group.

through across income groups depend on the extent to which low-income households purchase lower priced and lower margin products than high-income households. For example, there are only minor differences in the average prices paid by households for flour products, and thus differences in the long-run log pass-through of commodity costs to flour prices paid by different income groups is relatively small. On the other hand, the unit price paid by households in the lowest income quintile for coffee products is nearly 30 percent lower than that paid by households in the highest income quintile, and accordingly there are larger differences in the log pass-through of coffee commodity costs to prices across income groups.

6.2 Food-at-Home Inflation Inequality

While the previous analyses demonstrate how complete pass-through in levels can generate inflation inequality within narrow product categories, one may wonder whether complete pass-through in levels can generate quantitatively important differences in the overall inflation rates experienced by households over the income distribution. To this end, I now extend the analysis to overall food-at-home inflation.

When looking at overall food-at-home inflation, it is no longer possible to construct detailed matches from commodity costs to retail prices and measure pass-through in levels. Hence, in this section, I rely on log pass-through from two upstream price indices—producer price indices for Farm Products and Food Manufacturing—to downstream prices

faced by consumers (see Appendix Figure A14 for the evolution of these two producer price indices since 2006).

The analysis proceeds in three steps. First, I describe how I build a retail scanner price index that co-moves closely with food-at-home consumer price index from the Bureau of Labor Statistics. Second, I disaggregate this retail scanner price index by product "quality" (measured as unit price relative to other products in the same category) and show that price indices constructed from lower quality products have higher log pass-through of upstream producer price indices. Finally, I use the consumer panel data to construct food-at-home indices by income quintile and show that food-at-home inflation for the lowest quintile is more sensitive to upstream producer price indices and more volatile.

Reconstructing the food-at-home price index using scanner data. Beraja et al. (2019) show that price indices constructed from Nielsen Retail Scanner data can closely match consumer price indices released by the Bureau of Labor Statistics. I undertake a similar exercise as Beraja et al. (2019) to create a Retail Scanner Price Index that mimics the BLS food-at-home consumer price index. In particular, for all food products in the Nielsen data,[42] I calculate the inflation rate over the next year as the sales-weighted inflation rate for all products,

$$\pi_t^{\text{Retail Scanner Index}} = \frac{\sum_i \lambda_{i,t}(p_{i,t+4}/p_{i,t} - 1)}{\sum_i \lambda_{i,t}}, \tag{12}$$

where $p_{i,t+4}/p_{i,t} - 1$ is the year-over-year growth in the quantity-weighted average price of product i from quarter t to quarter $t + 4$ and $\lambda_{i,t}$ is the total sales of retailer-UPC i in quarter t.[43] I construct this Retail Scanner Price Index at two levels of disaggregation: the first takes each UPC as a product, and the second takes a unique retailer-UPC pair as a single product. The advantage of using the latter, finer level of disaggregation is that the same UPC may be priced quite differently across retail chains, but tends to have uniform pricing with a retailer (DellaVigna and Gentzkow 2019). On the other hand, taking the UPC as the lowest level of disaggregation increases the likelihood that the price of the product will be observed the following year, thus reducing the potential bias associated

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[42]I.e., products in dry grocery (department code 1), frozen foods (dept. 2), dairy (dept. 3), deli (dept. 4), packaged meat (dept. 5), and fresh produce (dept. 6).

[43]One difference between the retail scanner price index I construct and the BLS's consumer price index is that the BLS instead takes a quantity-weighted average of products within each category and then aggregates across categories using expenditure weights, while I take an expenditure-weighted average across all products. Using quantity weights to aggregate products within categories produces very similar results to those reported here. An advantage to using expenditure weights within product category is that it allows me to disaggregate the price index into "quality groups" using groups with equal sales within each category.

with discontinued products.[44] Nevertheless, we will see that price indices constructed at both levels of disaggregation produce similar results.

Appendix Figure A15 shows that year-over-year inflation for the resulting Retail Scanner Price Index—using either UPCs or retailer-UPC pairs as the lowest level of aggregation—co-moves closely with the BLS food-at-home consumer price index. The correlation between year-over-year inflation for the Retail Scanner Price Index (constructed at either level of aggregation) and the food-at-home CPI is 0.96.

Disaggregating by quality.   We can use the Retail Scanner Price Index to investigate how log pass-through of upstream prices, such as producer prices for Farm Products and Food Manufacturing, varies across products with different unit prices. To do so, in each quarter, I rank all products within each product category by average unit price over the prior year. For simplicity, I will refer to average unit price as "quality," referring to products within a category with the highest unit prices as the highest quality group.

For each product category in each quarter, I split products by quality into N groups with equal sales and construct a price index for products (across categories) in each quality group. For example, the price index for the highest-quality group reflects changes in prices of the highest-quality products across all product categories. Note that each quality group has an identical composition of sales across product categories.

Figure 10 plots year-over-year inflation rates for $N = 3$ quality groups. Two patterns emerge. First, inflation rates for the lowest-quality products are higher than inflation rates for the highest-quality products in each period. This secular gap in inflation rates is documented by Jaravel (2019), who attributes lower inflation rates for high-priced products to higher entry and competitive pressure. Second, the volatility of inflation rates for the lowest-quality products is greater than that for the highest-quality products. When average inflation is low—such as in 2009 and late 2015—the gap in inflation rates between low-quality and high-quality products shrinks to about 1pp. On the other hand, when average inflation is high—in late 2007 and 2010—the gap in inflation rates between low- and high-quality products widens to over 4pp. This latter pattern is novel and is consistent with complete pass-through in levels leading to higher sensitivity of inflation rates to upstream prices for low-price products.

I test the role of differential log pass-through in generating these differences in inflation

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[44]Appendix Table A10 reports the share of product observations for which a price in the following year is available at both levels of disaggregation.

Figure 10: Retail scanner price inflation for $N = 3$ groups of products split by unit price.
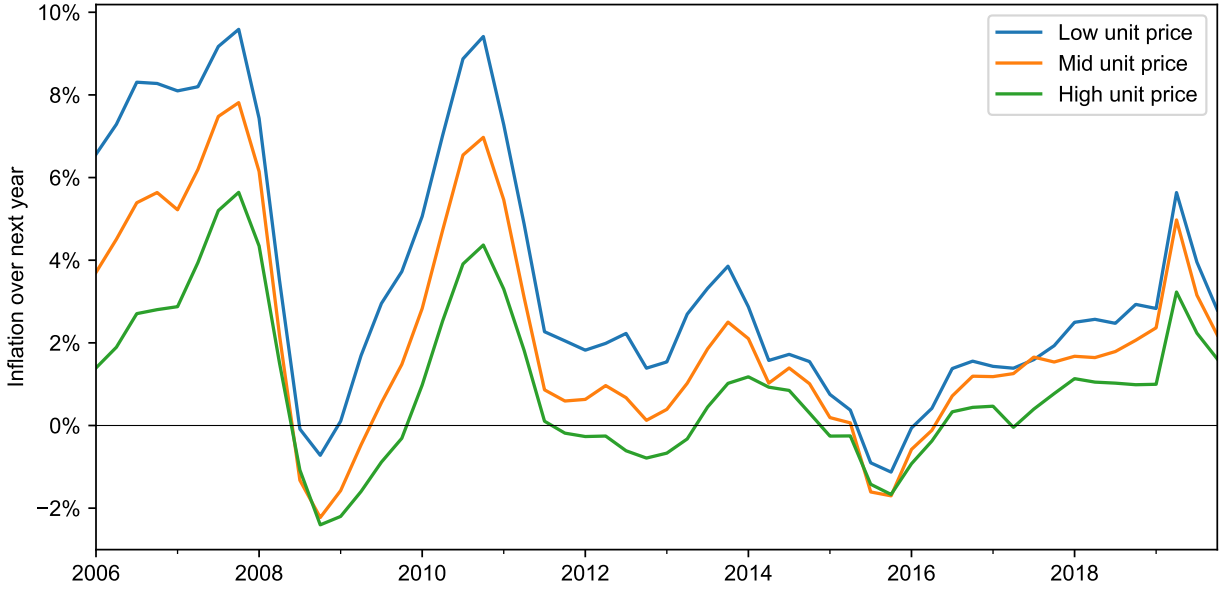

volatility using the distributed lag specification,

$$\Delta \log p_t^{\text{Retail Scanner Index},q} = a^q + \sum_{k=0}^{K} b_k^q \Delta \log \text{PPI}_{t-k} + \varepsilon_t, \tag{13}$$

where $\Delta \log p^{\text{Retail Scanner Index},q}$ is the log change in the price index for products in quality group q, $\Delta \log \text{PPI}_t$ is the log change in the upstream PPI (either the Farm Products PPI or the Food Manufacturing PPI), and $\sum_{k=0}^{K} b_k^q$ measures the long-run log pass-through of upstream PPI changes to retail price changes for quality group q. For consistency with the previous measures of pass-through for food products, I choose a horizon of one year (i.e., $K = 3$ quarters).

Figure 11 shows that estimates of long-run pass-through of Food Manufacturing and Farm Products PPIs to retail prices decline with product quality over $N = 10$ quality groups. The magnitudes of this decline are large: the log pass-through of Food Manufacturing price changes to products in the lowest quality decile is 0.75, compared to 0.39 for products in the highest-quality decile. These differences in log pass-through are independent of secular trends (i.e., unrelated to upstream PPIs) that cause inflation differences across unit price groups, which are instead captured by the intercept coefficients a^q.[45]

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[45] Appendix Figure A16 shows that the estimated intercept coefficients also decrease with quality, consistent with the secular drivers of inflation differences documented by Jaravel (2019).

Figure 11: Low unit price products have higher log pass-through of upstream prices.
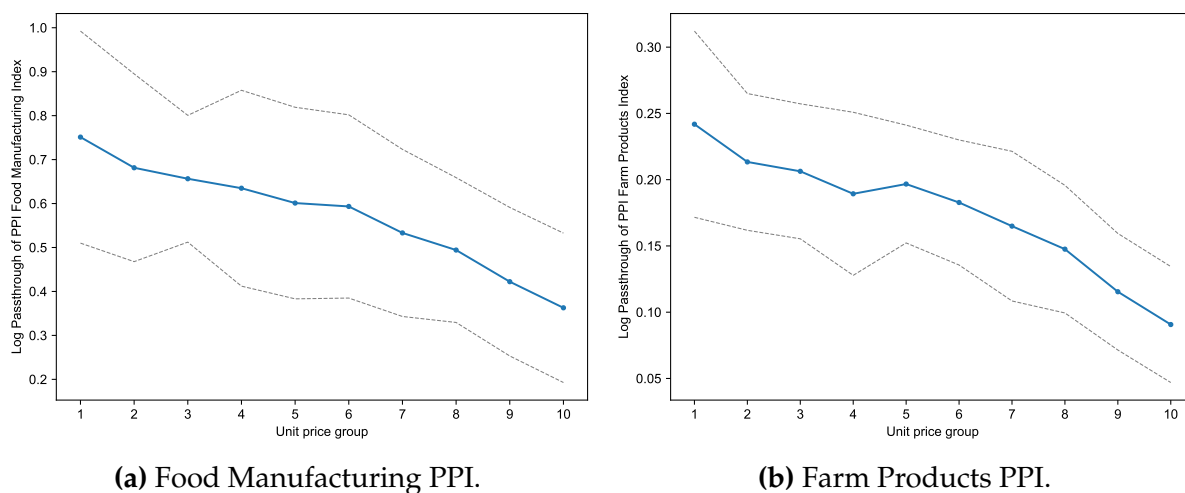
(a) Food Manufacturing PPI. **(b)** Farm Products PPI.

Note: Dotted lines indicate 95 percent confidence intervals using Driscoll-Kraay standard errors.

Differences across income groups.   How large are these effects for differences in food-at-home inflation experienced across income groups? I construct income group retail scanner price indices using the same approach as in (12), but now using expenditures from Nielsen Homescan panelists in each income group to weight price changes rather than sales in the scanner data.[46] As above, I construct these price indices using either each UPC or each retailer-UPC pair as the finest level of disaggregation.

Figure 12 plots the gap in food-at-home inflation rates experienced by the lowest and highest income quintiles since 2006. As documented by Jaravel (2019), this gap tends to be positive. However, there is significant cyclical variation in the level of inflation inequality that coincides with the average level of food-at-home inflation. In particular, inflation inequality grows when overall price levels are rising.

Figure 13a shows that the long-run log pass-through of upstream producer prices (the Food Manufacturing PPI) to the price index experienced by the lowest income quintile is 10 percent higher than that of the highest income quintile. This heightened sensitivity to upstream prices also translates into more volatile food-at-home inflation rates. Figure 13b shows that the variance of food-at-home inflation rates for the lowest income quintile is 20 percent higher than that of the highest income quintile. Appendix Figure A18 reports similar results by income decile.

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[46]The Nielsen Homescan panel reports household income in sixteen brackets. To sort households into income quintiles, I rank households first by income bracket, and within income bracket by total expenditures divided by the square-root of household size. This adjustment for household size follows from the OECD Income Distribution Database and Handbury (2021).

Figure 12: Gap in food-at-home inflation rates: Households in lowest vs. highest income quintile.
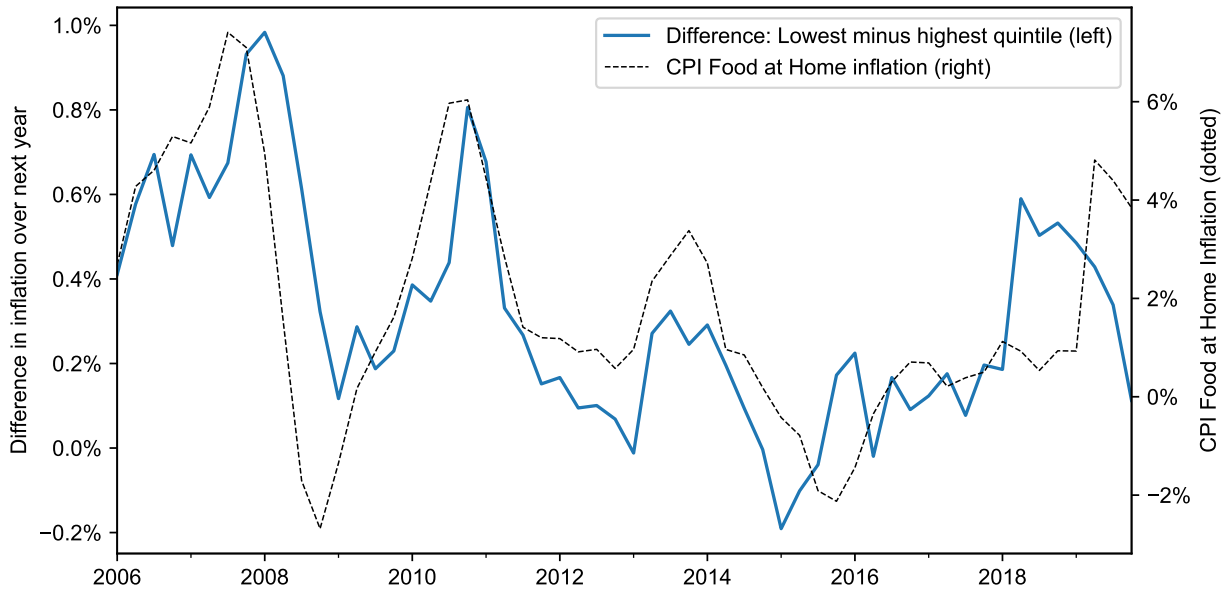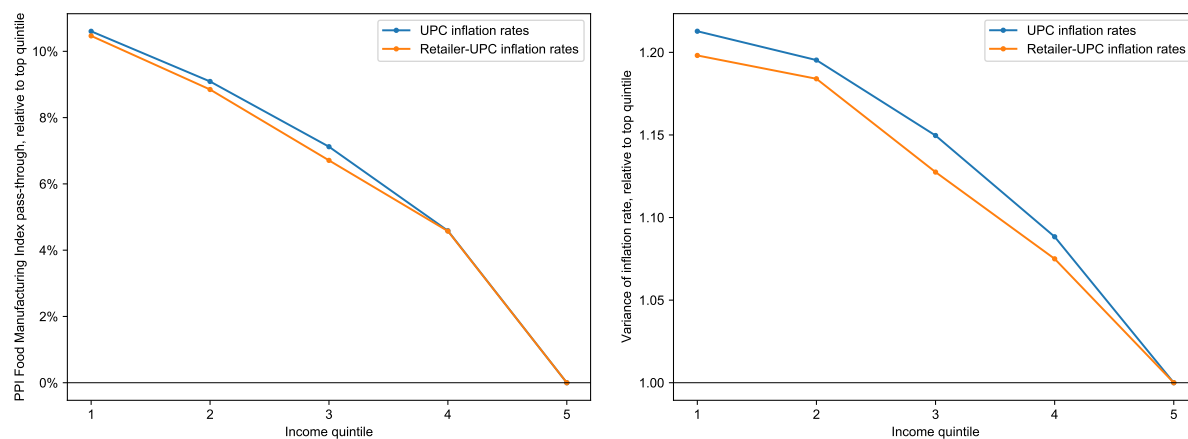

Figure 13: Food-at-home inflation for low-income more sensitive to upstream prices and higher variance.

(a) Pass-through of Food Manufacturing PPI.   **(b)** Variance of inflation rates.

Implications for food-at-home inflation, 2020–2023. At the time of writing, the most recent Nielsen data available ended in December 2020. In this section, I construct back-of-the-envelope estimates for inflation inequality over the recent period based on the evolution of upstream price indices. These estimates suggest sizable differences in inflation rates across income groups from January 2020 to January 2023, in part due to large increases in upstream costs over this period.

To estimate the price growth of a price index i from 2020 to 2023, I use fitted values for the intercept and long-run pass-through from (13) to calculate

$$\Delta \log(\text{PriceIndex}_{i,t}) \approx \underbrace{\alpha_i t}_{\text{Due to secular inflation rate}} + \underbrace{\rho_i^{\text{PPI}}(\Delta \log \text{PPI}_t)}_{\text{Due to pass-through}}, \tag{14}$$

where t is the number of quarters since January 2020, α_i is the intercept from the pass-through specification, and $\rho_i^{\text{PPI}} = \sum_{k=0}^{K} b_k^i$ is the long-run log pass-through of changes in the upstream PPI to price index i. The two terms in (14) capture two distinct channels that contribute to growth in price index i: secular trends in prices and how changes in upstream costs contribute to price growth.

Table 13 reports the predicted growth in price indices using this approach.[47] Due to both differences in pass-through and secular inflation rates, products in the lowest quality decile are predicted to have seen prices grow by 11pp more than products in the highest quality decile. Over 60 percent of the difference in price growth is due to differences in the pass-through of upstream costs. Similarly, the food-at-home price index for households in the lowest income quintile is predicted to have grown 15.6pp, compared to 13.7pp for households in the highest income quintile. Of the 1.8pp difference, two-thirds of the inflation inequality is due to differences in the pass-through of upstream producer prices, rather than secular differences in inflation rates.

7 Conclusion

Pass-through determines how changes in upstream costs are transmitted to downstream firms and consumers. This paper documents that observed patterns of incomplete log

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[47]These baseline estimates aggregate inflation rates at the UPC level and use changes in the Food Manufacturing PPI for upstream costs. Using retailer-UPC pairs as the lowest level of aggregation produces similar results, shown in Appendix Table A11. Appendix Table A11 also reports results using the Farm Products PPI. Long-run pass-through estimates from the Farm Products PPI to downstream price indices are noisier, attenuating the portion of the differences in predicted price growth due to differential pass-through, though the estimates are qualitatively similar.

Table 13: Unequal price growth from January 2020 to January 2023: Predicted changes using growth in upstream PPI (Food Manufacturing).

	Predicted growth in price index		
	Total	Due to PPI pass-through	Due to intercept
Products in unit price decile 1	20.7pp	16.0pp	4.7pp
Products in unit price decile 10	9.3pp	8.9pp	0.3pp
Difference	**11.4pp**	**7.1pp**	**4.3pp**
Lowest income quintile	15.6pp	12.7pp	2.8pp
Highest income quintile	13.7pp	11.5pp	2.2pp
Difference	**1.8pp**	**1.2pp**	**0.6pp**

pass-through may be better understood in terms of complete pass-through in levels and a lack of adjustment in unit margins. In the retail gasoline and food markets studied in this paper, complete pass-through in levels explains both the extent of pass-through and cross-sectional variation in log pass-through, and empirical exercises suggest similar patterns appear across a broader array of food-at-home products purchased by households.

Complete pass-through in levels suggests that considerations of price-setters beyond the elasticity of demand may be important to explain observed prices. For example, I present empirical evidence suggesting that the shape of firms' demand curves is not sufficient to explain patterns of pass-through, and that these price dynamics are better explained by a model in which overhead costs and risk shape the prices set by firms. The proposed model can also explain other features of the data, such as the lack of new firm entry when commodity costs rise.

Finally, this paper shows that complete pass-through in levels is useful in predicting pricing patterns, using the example of inflation inequality. For example, I show that this pricing behavior generates variation in inflation inequality over the commodity cost cycle, wherein rising commodity costs lead to surges in inflation inequality. These predictions are borne out in the data and constitute an important part of the gap between food-at-home inflation rates experienced by high- and low-income households from 2020–2023.

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Online Appendix

(Not for publication)

Appendix A   Additional Tables and Figures

Table A1: Unit root tests for commodity series.

	Levels			First differences		
	Autocorrelation coefficient (β)	Standard error	ADF test p-value	Autocorrelation coefficient (γ)	Standard error	ADF test p-value
Canada Crude*	0.982	(0.009)	0.721	-0.090	(0.097)	0.000
Canada Wholesale*	0.987	(0.010)	0.961	0.139	(0.048)	0.000
Australia Terminal Unleaded	0.996	(0.007)	0.731	0.449	(0.058)	0.000
Australia Terminal Premium Unleaded	0.995	(0.006)	0.665	0.442	(0.058)	0.000
Australia Terminal Diesel	0.999	(0.007)	0.919	0.302	(0.142)	0.000
Beef Farm Price	0.993	(0.007)	0.555	0.280	(0.041)	0.000
Pork Farm Price	0.930	(0.018)	0.000	0.170	(0.039)	0.000
Skim Milk*	0.950	(0.006)	0.498	-0.029	(0.018)	0.000
Butterfat*	0.904	(0.008)	0.149	0.150	(0.018)	0.000
Coffee	0.983	(0.010)	0.322	0.229	(0.052)	0.000
Sugar	0.975	(0.018)	0.242	0.199	(0.083)	0.000
Beef	0.997	(0.008)	0.939	0.238	(0.042)	0.000
Rice	0.987	(0.010)	0.165	0.347	(0.078)	0.000
Flour	0.984	(0.011)	0.343	0.213	(0.047)	0.000
Orange	0.967	(0.013)	0.028	0.238	(0.045)	0.000

Note: Columns 1 and 4 report coefficients estimated from the specifications,

$$c_t = \beta c_{t-1} + \varepsilon_t,$$
$$\Delta c_t = \gamma \Delta c_{t-1} + \hat{\varepsilon}_t.$$

Columns 2 and 5 report Newey-West standard errors with four lags. Columns 3 and 6 report the p-value from Augmented Dickey-Fuller tests for unit roots, where the null hypothesis is that the series is a unit root process. * Asterisk rows contain multiple series by markets. For these rows, standard errors are reported are Driscoll-Kraay standard errors, and the reported Augmented Dickey-Fuller test for unit root report the maximum p-value across all markets.

Table A2: Granger causality tests for commodity and retail prices.

| | Granger causality test p-value | |
	1 to 2	2 to 1
Canada, city-level, 2007–2022		
Crude to wholesale	0.003	0.908
Crude to retail (excl. taxes)	0.053	0.999
Wholesale to retail (excl. taxes)	0.000	1.000
Australia, station-level, 2001–2022		
Terminal ULP to Station Price ULP	0.000	0.001
Terminal PULP to Station Price PULP	0.000	0.001
Terminal Diesel to Station Price Diesel	0.000	0.120
USDA ERS		
Beef Farm to Wholesale	0.000	0.205
Beef Farm to Retail	0.000	0.126
Beef Farm to Fresh Retail	0.044	0.567
Beef Wholesale to Retail*	0.000	0.003
Beef Wholesale to Fresh Retail	0.000	0.441
Pork Farm to Wholesale*	0.000	0.007
Pork Farm to Retail	0.000	0.069
Pork Wholesale to Retail	0.063	0.785
Dairy Commodity to Whole Retail**	0.000	0.877
Dairy Commodity to Reduced Fat Retail**	0.003	0.826
U.S. CPI commodities		
Coffee Commodity (IMF) to Retail (CPI)**	0.000	0.334
Sugar Commodity (IMF) to Retail (CPI)**	0.003	0.652
Beef Commodity (IMF) to Retail (CPI)**	0.688	0.956
Rice Commodity (IMF) to Retail (CPI)**	0.353	0.877
Flour Commodity (IMF) to Retail (CPI)**	0.700	0.931
Orange Commodity (IMF) to Retail (CPI)**	0.053	0.979

Note: * Starred entries note relationships where reverse causality is a likely concern. ** Uses four lags instead of twelve

Table A3: IMF primary commodity prices and sources.

Commodity series	IMF Series ID	Description
Global price of Coffee, Other Mild Arabica	PCOFFOTMUSDM	Coffee, Other Mild Arabicas, International Coffee Organization New York cash price, ex-dock New York
Global price of Sugar, No. 16, US	PSUGAUSAUSDM	Sugar, U.S. import price, contract no. 16 futures position
Global price of Beef	PBEEFUSDM	Beef, Australian and New Zealand 85% lean fores, CIF U.S. import price
Global price of Rice, Thailand	PRICENPQUSDM	Rice, 5 percent broken milled white rice, Thailand nominal price quote
Global price of Wheat	PWHEAMTUSDM	Wheat, No. 1. Hard Red Winter, ordinary protein, Kansas City
Global price of Orange	PORANGUSDM	Generic 1st 'JO' Future

Table A4: Food products commodity and retail price series with unit conversion factors.

Commodity series	IMF Series ID	Units	BLS Average Price Data series	Series ID[48]	Unit conversion factor
Global price of Coffee, Other Mild Arabica	PCOFFOTMUSDM	Cents per Pound	Coffee, 100 percent, ground roast, per lb.	717311, 717312	1.235 (19% weight lost in roasting process[49])
Global price of Sugar, No. 16, US	PSUGAUSAUSDM	Cents per Pound	Sugar, white, per lb.	715211, 715212	1
Global price of Beef	PBEEFUSDM	Cents per Pound	Ground beef, 100% beef, per lb. (453.6 gm)	703112	1
Global price of Rice, Thailand	PRICENPQUSDM	Dollars per Metric Ton	Rice, white, long grain, uncooked, per lb. (453.6 gm)	701312	0.0454 (100 dollars per cent / 2204.62 lbs per metric ton)
Global price of Wheat	PWHEAMTUSDM	Dollars per Metric Ton	Flour, white, all purpose, per lb. (453.6 gm)	701111	0.0650 (100 dollars per cent / 2204.62 lbs per metric ton wheat / 42 lbs flour per 60 lbs (1 bushel) wheat)
Global price of Orange	PORANGUSDM	Dollars per Pound	Orange juice, frozen concentrate, 12 oz. can, per 16 oz. (473.2 mL)	713111	66.9 (100 dollars per cent × 41.8 retail brix / 62.5 futures brix[50])

[48] For some products, multiple series are available which track different package sizes.
[49] Nakamura and Zerom (2010).
[50] Dutta et al. (2002) for retail brix content and ICE for futures contract brix content.

Table A5: Higher-priced products exhibit lower log pass-through, with no systematic difference in level pass-through: Five groups.

Panel A: In percentages

	Retail price inflation		
	Rice	Flour	Coffee
Commodity Inflation	0.248**	0.077**	0.125**
	(0.017)	(0.008)	(0.013)
Commodity Inflation × Unit Price Group 2	−0.070**	−0.003	−0.034
	(0.017)	(0.018)	(0.022)
Commodity Inflation × Unit Price Group 3	−0.095**	−0.004	−0.089**
	(0.015)	(0.005)	(0.021)
Commodity Inflation × Unit Price Group 4	−0.127**	−0.045**	−0.102**
	(0.018)	(0.010)	(0.019)
Commodity Inflation × Unit Price Group 5	−0.197**	−0.055**	−0.106**
	(0.021)	(0.008)	(0.015)
UPC FEs	Yes	Yes	Yes
N (thousands)	399.4	101.4	1570.0
R^2	0.16	0.06	0.15

Panel B: In levels

	Δ Retail price		
	Rice	Flour	Coffee
Δ Commodity Price	0.056**	0.033**	0.061**
	(0.006)	(0.006)	(0.007)
Δ Commodity Price × Unit Price Group 2	0.001	0.013	−0.001
	(0.008)	(0.009)	(0.012)
Δ Commodity Price × Unit Price Group 3	0.010	0.016**	−0.030
	(0.007)	(0.008)	(0.018)
Δ Commodity Price × Unit Price Group 4	0.006	−0.017	−0.035*
	(0.008)	(0.021)	(0.021)
Δ Commodity Price × Unit Price Group 5	0.006	−0.029**	−0.027**
	(0.016)	(0.013)	(0.014)
UPC FEs	Yes	Yes	Yes
N (thousands)	399.4	101.4	1570.0
R^2	0.07	0.05	0.17

Note: Panel A reports results from specification (5), and panel B reports results from specification (6). The three columns use products from rice, all purpose white flour, and roasted coffee, respectively. In each quarter, each retailer-UPC pair is assigned to five groups with equal sales by (quantity-weighted) average unit price over the past year. Unit Price Groups 2–5 are indicators for whether a retailer-UPC pair is assigned to the low-mid to the highest unit price group. Standard errors clustered by brand. * indicates significance at 10%, ** at 5%.

Table A6: Comparison to two alternative models of log pass-through: Rice products.

Retail product inflation	(1)	Market share (2)	(3)	Buyer elasticity (4)	(5)
Commodity Inflation	0.118**	0.174**	0.113**	0.133**	0.117**
	(0.006)	(0.015)	(0.010)	(0.016)	(0.008)
Commodity Infl. × Log(Unit Price)	−0.100**		−0.101**		−0.105**
	(0.012)		(0.012)		(0.016)
Commodity Infl. × Log(Brand Sales Share)		0.006*	−0.002		
		(0.003)	(0.003)		
Wage Inflation				1.179**	0.452**
				(0.105)	(0.113)
Wage Infl. × Log(Buyer Income)				−2.364**	−0.654
				(0.770)	(0.566)
UPC FEs	Yes	Yes	Yes	Yes	Yes
N (thousands)	399.4	399.4	399.4	329.8	329.8
R^2	0.18	0.12	0.18	0.13	0.19

Note: Log(Unit Price), Log(Brand Sales Share), and Log(Buyer Income) are all normalized relative to the average within each quarter, so that these three terms represent log deviations from the average unit price, sales share, and buyer income across all products in the quarter. Standard errors clustered by brand. * indicates significance at 10%, ** at 5%.

Table A7: Comparison to two alternative models of log pass-through: Flour products.

Retail product inflation	(1)	Market share		Buyer elasticity	
		(2)	(3)	(4)	(5)
Commodity Inflation	0.053**	0.049**	0.042**	0.042**	0.047**
	(0.006)	(0.014)	(0.007)	(0.009)	(0.008)
Commodity Infl. × Log(Unit Price)	−0.065**		−0.069**		−0.075**
	(0.010)		(0.010)		(0.019)
Commodity Infl. × Log(Brand Sales Share)		−0.003	−0.005		
		(0.005)	(0.003)		
Wage Inflation				2.145**	1.791**
				(0.280)	(0.268)
Wage Infl. × Log(Buyer Income)				0.222	0.553
				(2.487)	(2.530)
UPC FEs	Yes	Yes	Yes	Yes	Yes
N (thousands)	101.4	101.4	101.4	79.6	79.6
R^2	0.04	0.03	0.04	0.04	0.05

Note: Log(Unit Price), Log(Brand Sales Share), and Log(Buyer Income) are all normalized relative to the average within each quarter, so that these three terms represent log deviations from the average unit price, sales share, and buyer income across all products in the quarter. Standard errors clustered by brand. * indicates significance at 10%, ** at 5%.

Table A8: Comparison to two alternative models of log pass-through: Coffee products.

		Market share		Buyer elasticity	
Retail product inflation	(1)	(2)	(3)	(4)	(5)
Commodity Inflation	0.051**	0.141**	0.112**	0.058**	0.056**
	(0.010)	(0.029)	(0.018)	(0.021)	(0.011)
Commodity Infl. × Log(Unit Price)	−0.080**		−0.060**		−0.085**
	(0.014)		(0.015)		(0.016)
Commodity Infl. × Log(Brand Sales Share)		0.024**	0.017**		
		(0.005)	(0.004)		
Wage Inflation				−0.240	−0.794**
				(0.245)	(0.264)
Wage Infl. × Log(Buyer Income)				0.598	2.173*
				(1.127)	(1.167)
UPC FEs	Yes	Yes	Yes	Yes	Yes
N (thousands)	1570.0	1570.0	1570.0	1269.3	1269.3
R^2	0.16	0.14	0.16	0.13	0.16

Note: Log(Unit Price), Log(Brand Sales Share), and Log(Buyer Income) are all normalized relative to the average within each quarter, so that these three terms represent log deviations from the average unit price, sales share, and buyer income across all products in the quarter. Standard errors clustered by brand. * indicates significance at 10%, ** at 5%.

Table A9: Exploiting variation in margins across retailers: Summary of results across all product modules using leave-one-out change in prices.

Share of modules	Unweighted	Observations-weighted	Sales-weighted
Panel A: In levels			
Positive coefficient	9.1	7.7	9.1
Negative coefficient	19.5	32.8	33.5
Panel B: In logs			
Positive coefficient	1.8	0.8	0.4
Negative coefficient	56.8	79.3	79.5

Note: Summary of results from specifications (7) (for panel A) and (8) (for panel B) estimated across 616 product modules. Each cell reports the fraction of product modules for which the estimated interaction between the average UPC price change (in levels or logs) and the relative price of the product at the retailer is significant at a 5% level. Driscoll-Kraay standard errors used in all specifications.

Table A10: Percent of expenditures matched to retail scanner and inflation data, by income group.

Income quintile	Matched to UPC		Matched to retailer-UPC	
	Total	With infl.	Total	With infl.
1	60.2	52.7	22.5	18.5
2	59.9	52.6	23.1	19.0
3	60.2	53.5	24.0	20.1
4	60.7	54.5	25.7	21.7
5	59.7	52.6	27.2	22.7

Table A11: Unequal price growth from January 2020 to January 2023: Predicted changes using alternative measures.

		Predicted growth in price index	
	Total	Due to PPI pass-through	Due to intercept
Food Manufacturing PPI, UPC aggregation (baseline)			
Products in quality decile 1	20.7pp	16.0pp	4.7pp
Products in quality decile 10	9.3pp	8.9pp	0.3pp
Difference	**11.4pp**	**7.1pp**	**4.3pp**
Lowest income quintile	15.6pp	12.7pp	2.8pp
Highest income quintile	13.7pp	11.5pp	2.2pp
Difference	**1.8pp**	**1.2pp**	**0.6pp**
Food Manufacturing PPI, Retailer-UPC aggregation			
Quality decile 1	26.3pp	19.5pp	6.9pp
Quality decile 10	7.5pp	9.4pp	-1.9pp
Difference	**18.9pp**	**10.1pp**	**8.8pp**
Lowest income quintile	18.7pp	13.5pp	5.2pp
Highest income quintile	16.3pp	12.3pp	4.0pp
Difference	**2.4pp**	**1.3pp**	**1.1pp**
Farm Products PPI, UPC aggregation			
Quality decile 1	18.1pp	10.2pp	7.9pp
Quality decile 10	7.0pp	4.9pp	2.2pp
Difference	**11.0pp**	**5.3pp**	**5.7pp**
Lowest income quintile	12.7pp	7.2pp	5.5pp
Highest income quintile	11.5pp	6.9pp	4.6pp
Difference	**1.3pp**	**0.3pp**	**0.9pp**
Farm Products PPI, Retailer-UPC aggregation			
Quality decile 1	23.8pp	13.0pp	10.8pp
Quality decile 10	4.9pp	4.9pp	-0.0pp
Difference	**18.9pp**	**8.1pp**	**10.8pp**
Lowest income quintile	15.7pp	7.5pp	8.1pp
Highest income quintile	13.8pp	7.2pp	6.6pp
Difference	**1.9pp**	**0.4pp**	**1.5pp**

Figure A1: Retail unleaded petrol (ULP) price and terminal gas price for Rottnest Island Authority station at Thompson Bay Fuel Jetty, Cockburn, Perth, Australia.

Figure A2: Premium unleaded petrol price (PULP) pass-through in levels (top) and in logs (bottom).
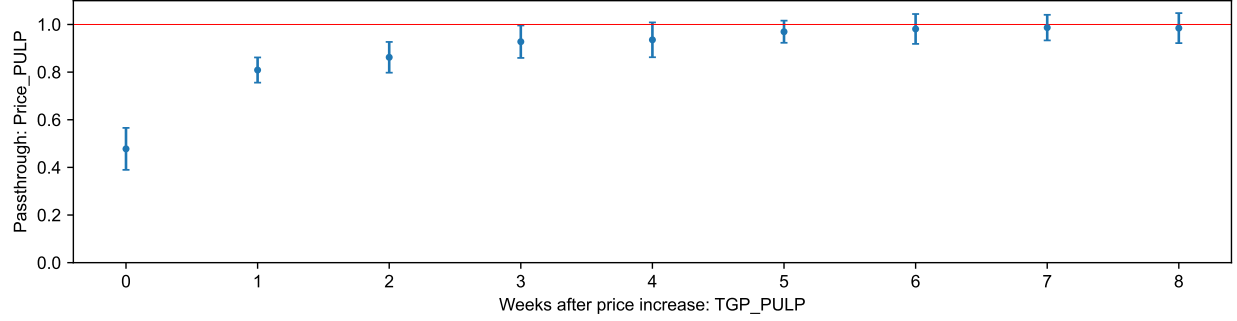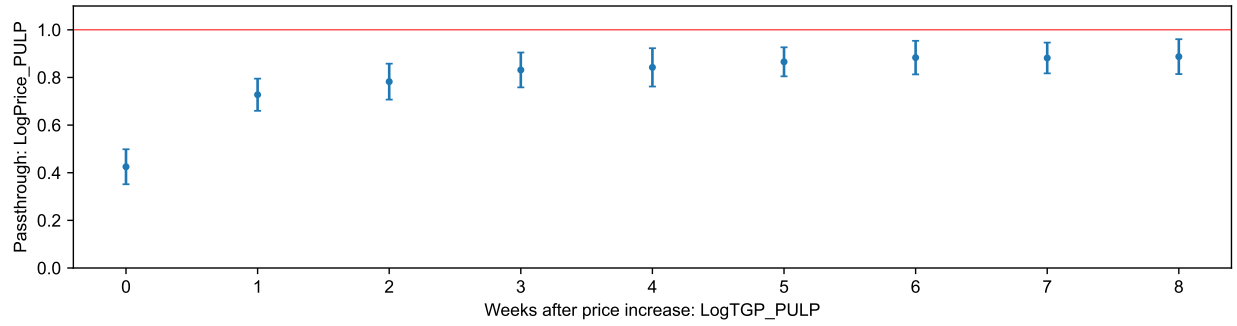
(a) Pass-through in levels.

(b) Pass-through in logs.

Note: Panels (a) and (b) show cumulative pass-through estimated from the specifications,

$$\Delta p_{i,t} = \sum_{k=0}^{k=8} b_k \Delta c_{i,t-k} + a_i + \varepsilon_{i,t}.$$

$$\Delta \log p_{i,t} = \sum_{k=0}^{k=8} \beta_k \Delta \log c_{i,t-k} + \alpha_i + \varepsilon_{i,t}.$$

Standard errors are two-way clustered by postcode and year (Driscoll-Kraay panel standard errors are similar), and standard errors for cumulative pass-through coefficients $\sum_{k=0}^{t} b_k$ and $\sum_{k=0}^{t} \beta_k$ are computed using the delta method.

Figure A3: Comovement of retail gas margins with strength of weekly price cycles.

(a) Unleaded petrol (ULP).　　　　**(b)** Premium unleaded petrol (PULP).
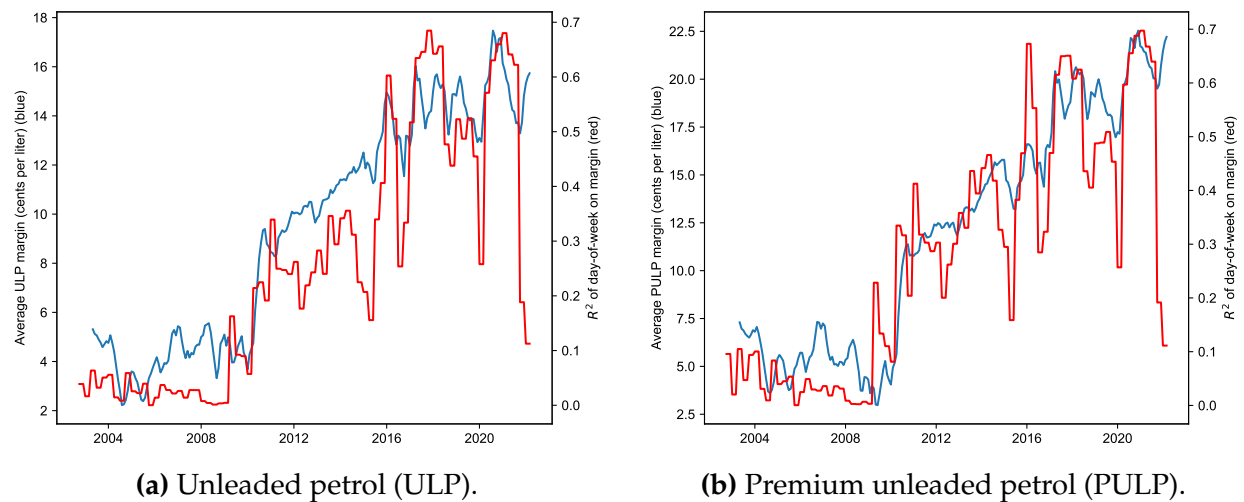
Note: In each panel, the blue line (left axis) plots the six-month moving average of margins across all stations. The red line (right axis) plots the R^2 from a regression of gas station margins of day-of-week dummies for each quarter.

Figure A4: Pass-through in levels across groups of relative price.

(a) Unleaded petrol:
Groups within neighborhood.

(b) Unleaded petrol:
Groups within postcode.

(c) Unleaded petrol:
Groups across all stations.

(d) Premium unleaded petrol:
Groups within neighborhood.

(e) Premium unleaded petrol:
Groups within postcode.

(f) Premium unleaded petrol:
Groups across all stations.
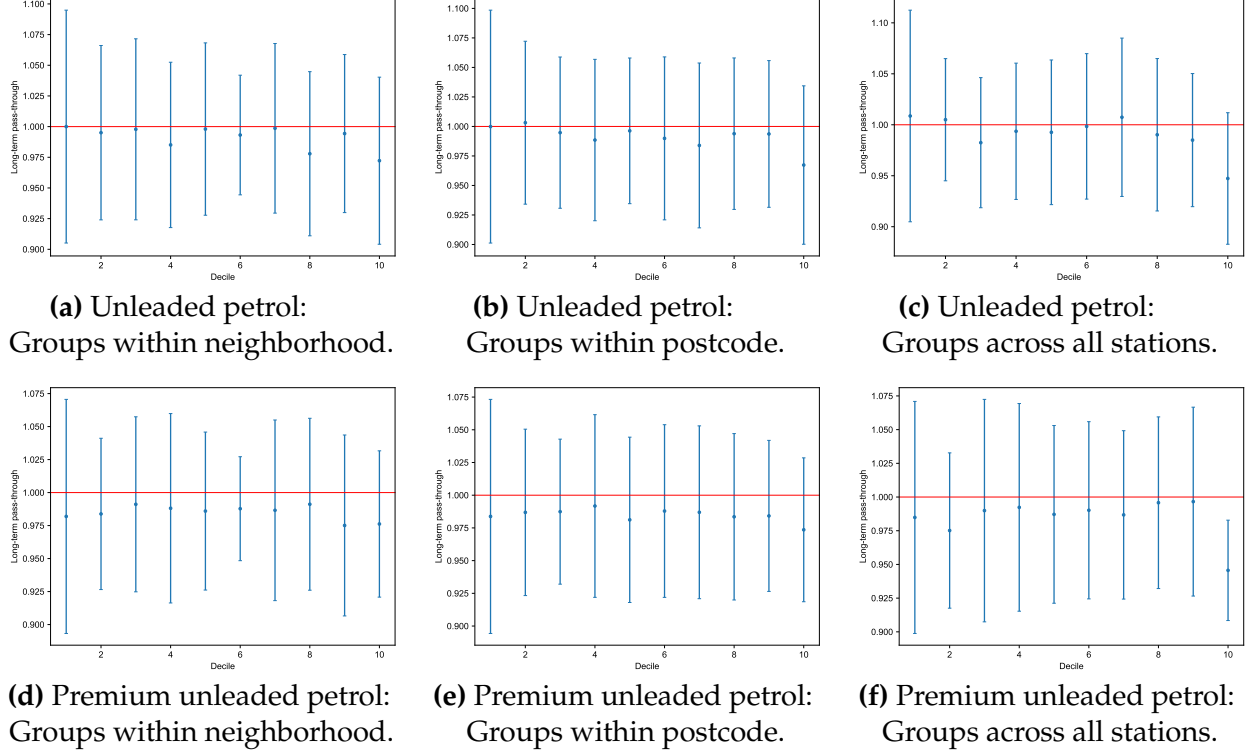
Note: These charts plot the long-run pass-through estimated from the specification,

$$\Delta p_{i,t} = \sum_{k=0}^{k=8} b_k \Delta c_{i,t-k} + a_i + \varepsilon_{i,t}.$$

Standard errors are two-way clustered by postcode and year (Driscoll-Kraay panel standard errors are similar), and standard errors for cumulative pass-through coefficients $\sum_{k=0}^{t} b_k$ and $\sum_{k=0}^{t} \beta_k$ are computed using the delta method. For each figure, the specification is estimated separately across ten deciles of RelativePrice$_i$, where

$$\text{RelativePrice}_i = \frac{1}{T_i} \sum_t \left(\text{Price}_{i,t} - \frac{1}{|N_t(i)|} \sum_{j \in N_t(i)} \text{Price}_{j,t} \right),$$

T_i is the number of daily observations in the sample for station i, Price$_{i,t}$ is i's retail price on day t, and $N_t(i)$ is the set of stations in i's neighborhood or postcode on date t.

Figure A5: Price cycles in unleaded petrol (ULP) for BP station at 549 Abernethy Rd, Kewdale, Perth, Australia.
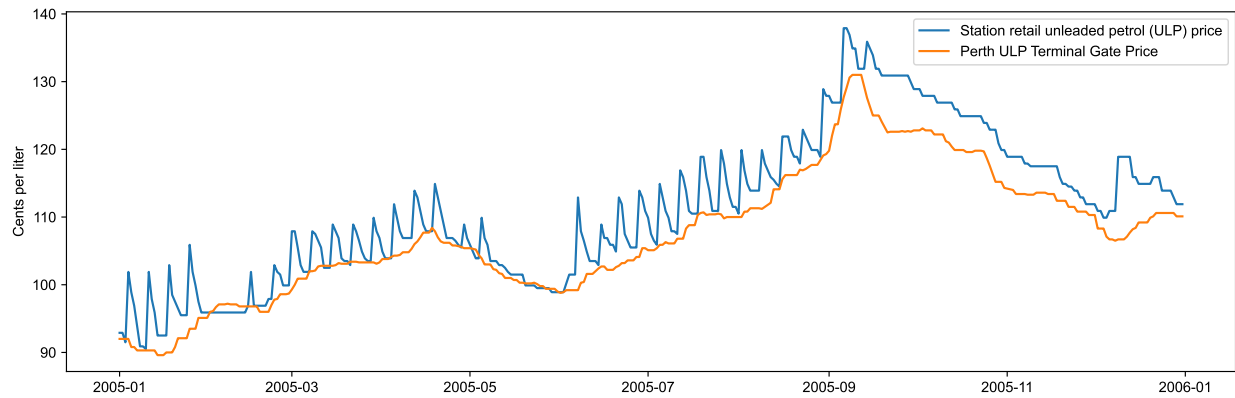
(a) 2005.
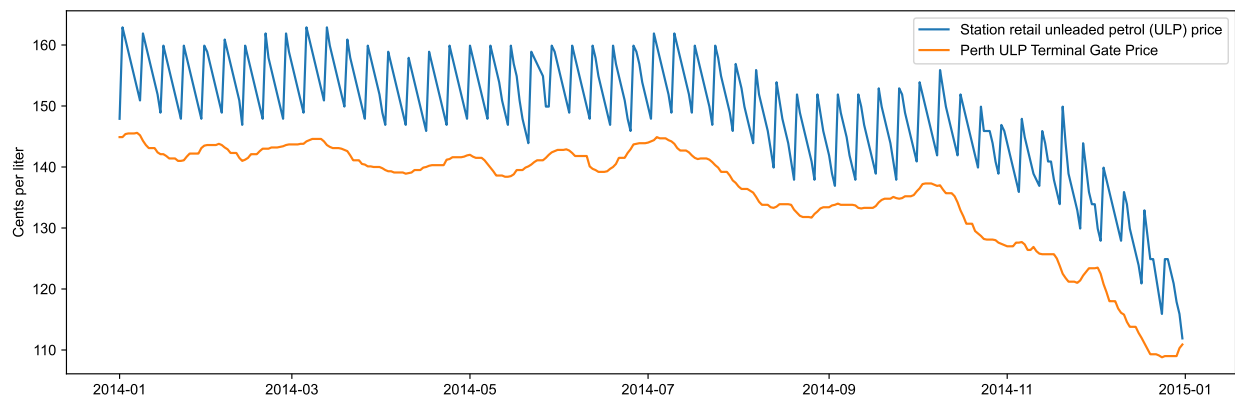
(b) 2014.
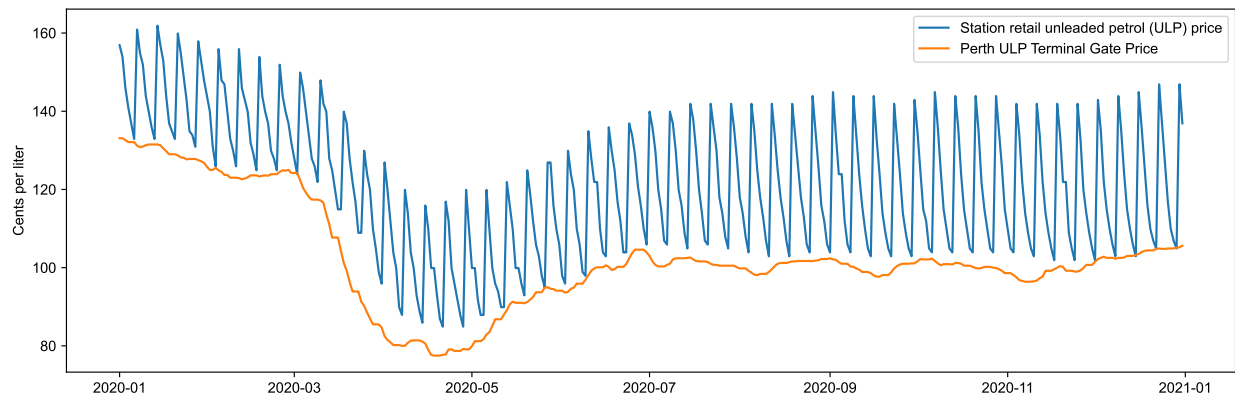
(c) 2020.

Figure A6: Pass-through of U.S. refiner wholesale prices to retail prices, 1983–2021.

(a) Logs ($\beta = 0.74$).

(b) Levels ($\beta = 0.99$).
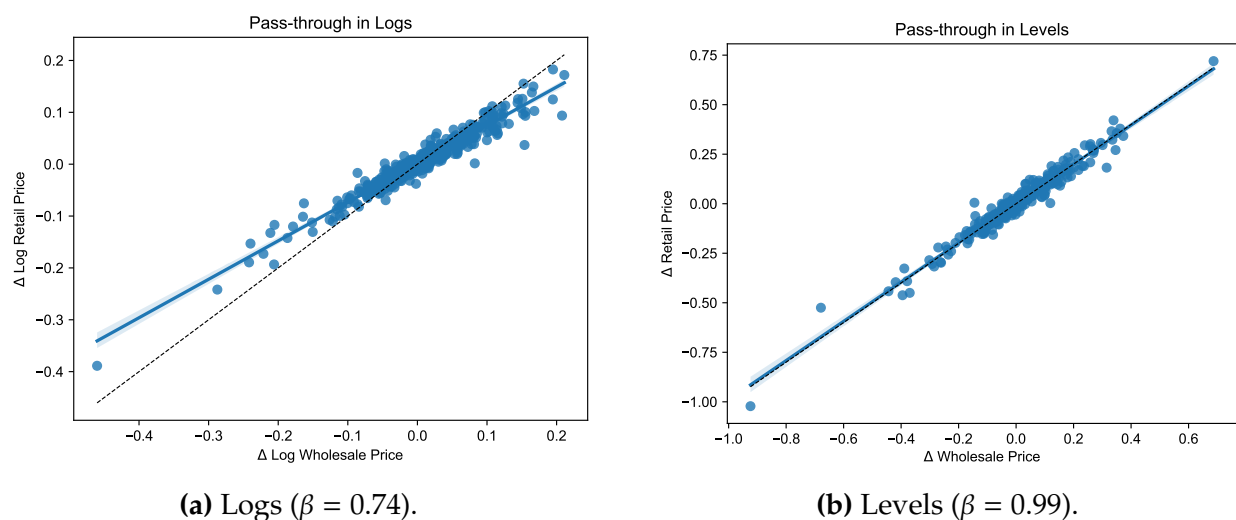
Note: Refiner wholesale prices are monthly U.S. refiner gasoline prices for sale through retail outlets from the Energy Information Administration (EIA). Retail gasoline prices are from BLS Average Price Data. Price changes measured over one month.

Figure A7: Passthrough of sugar commodity costs to retail prices.
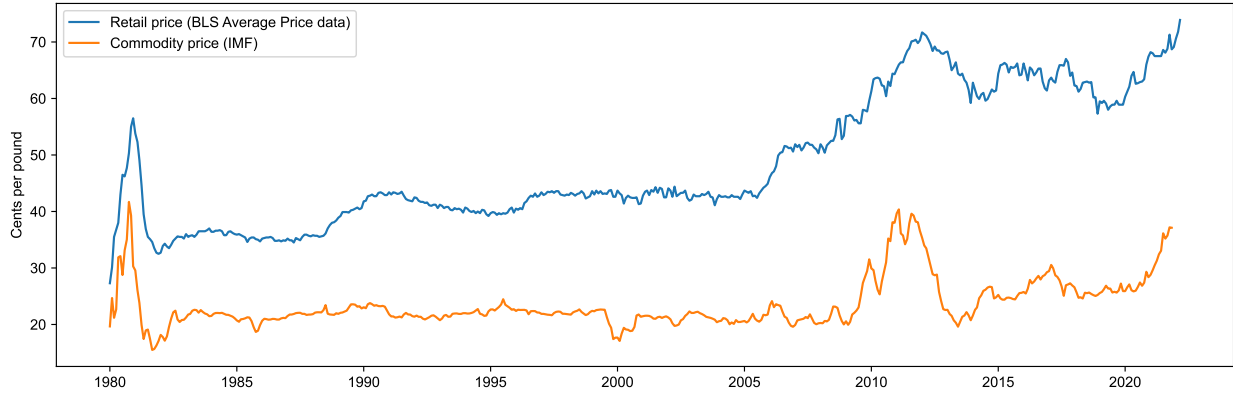
(a) Sugar No. 16 commodity costs (IMF) and retail white granulated sugar prices (U.S. CPI).
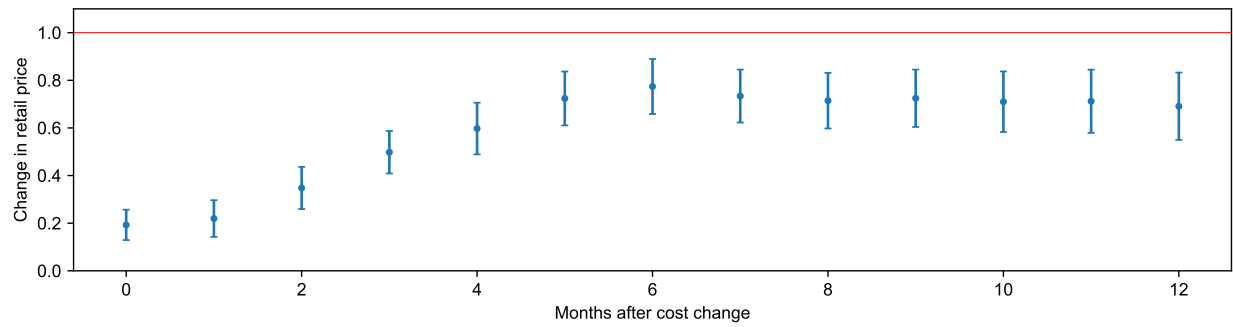
(b) Pass-through in levels.
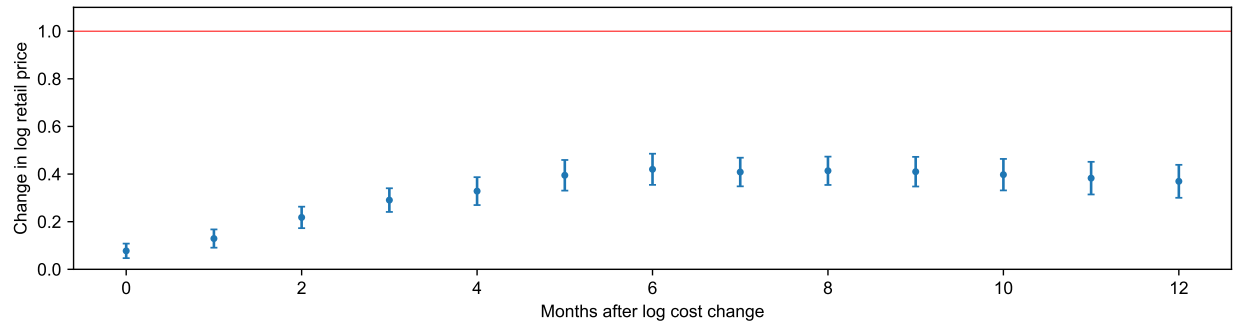
(c) Pass-through in logs.

Note: Panel (a) plots the time series of the commodity price from the IMF and the Average Price Data series from the BLS. The series are adjusted by the conversion factors in Appendix Table A4 so that the two series are in comparable units. Panels (b) and (c) plot the cumulative pass-through to month T, $\sum_{k=0}^{T} b_k$, from the specifications (2) and (3), using a total horizon of $K = 12$ months.
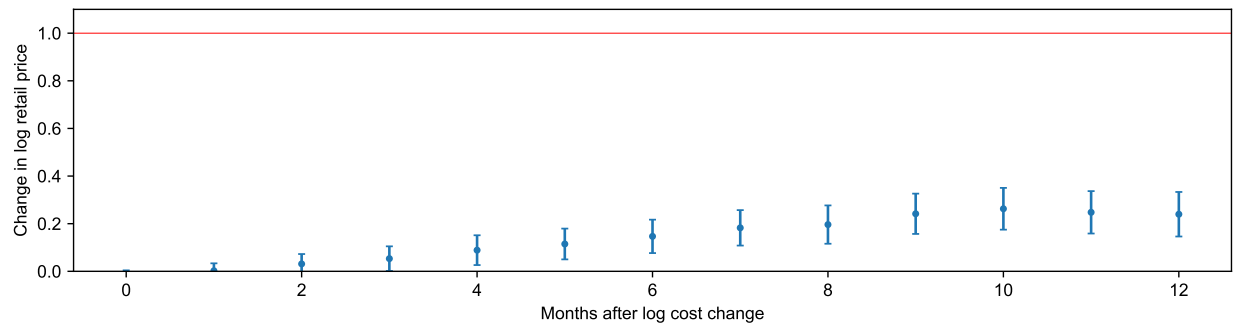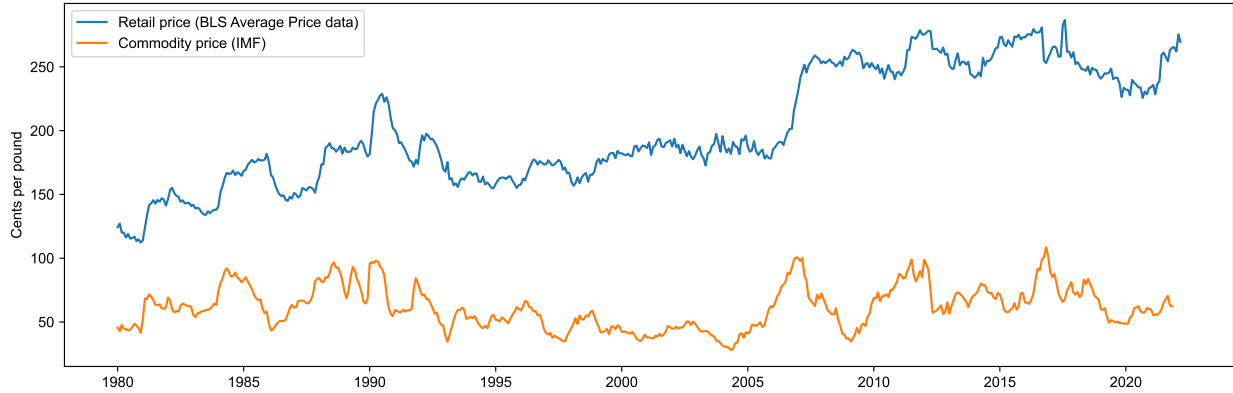
Figure A8: Passthrough of beef commodity costs to retail prices.
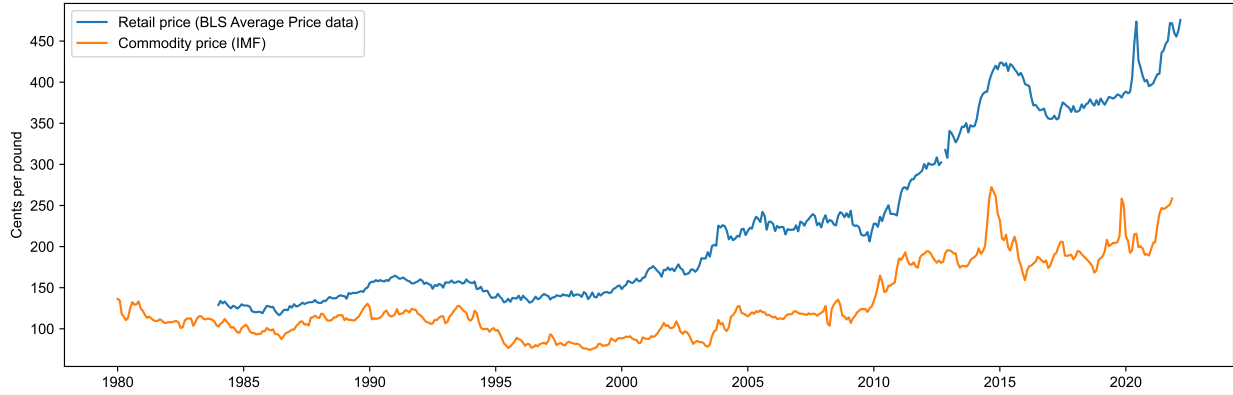
(a) Beef commodity costs (IMF) and retail ground beef prices (U.S. CPI).
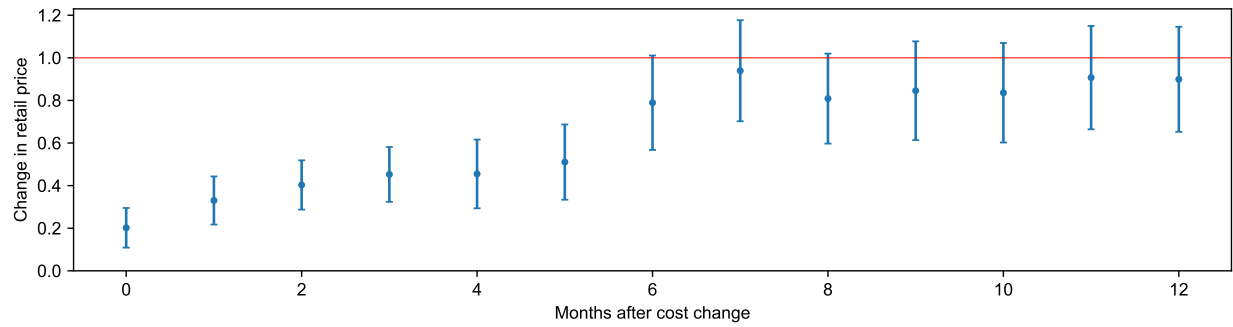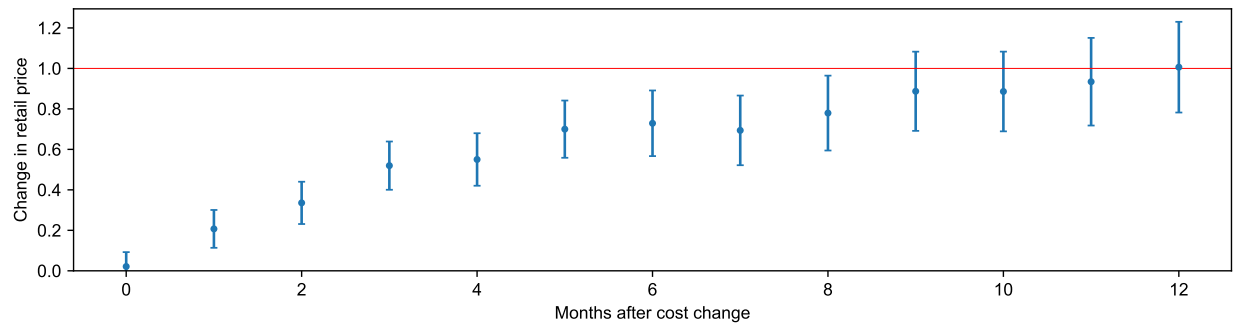

(b) Pass-through in levels.
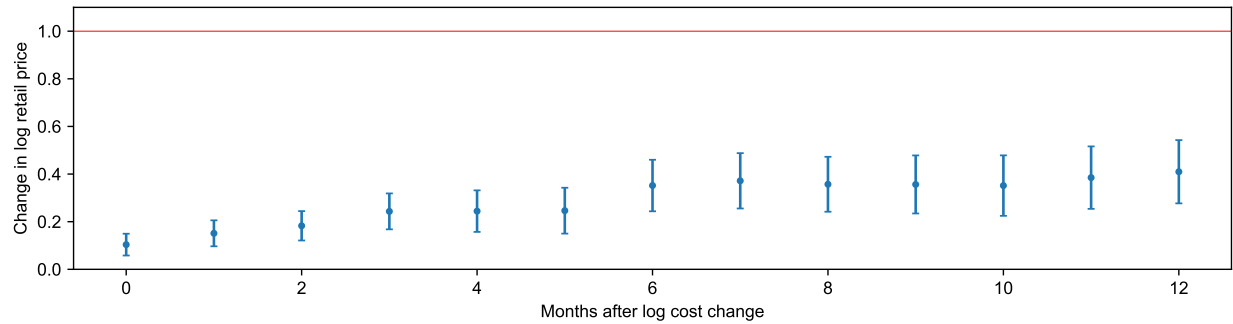
(c) Pass-through in logs.

Note: Panel (a) plots the time series of the commodity price from the IMF and the Average Price Data series from the BLS. The series are adjusted by the conversion factors in Appendix Table A4 so that the two series are in comparable units. Panels (b) and (c) plot the cumulative pass-through to month T, $\sum_{k=0}^{T} b_k$, from the specifications (2) and (3), using a total horizon of $K = 12$ months.

Figure A9: Passthrough of rice commodity costs to retail prices.
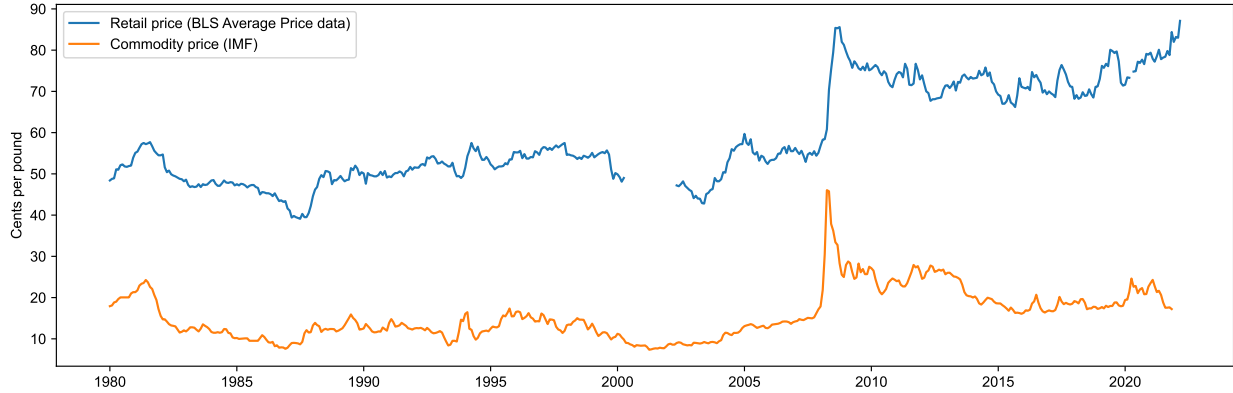
(a) Thailand rice commodity costs (IMF) and retail long-grain white rice prices (U.S. CPI).
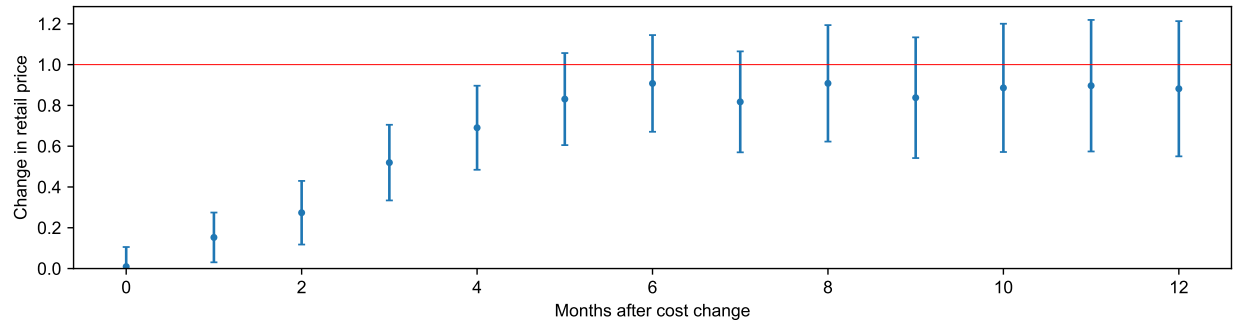

(b) Pass-through in levels.
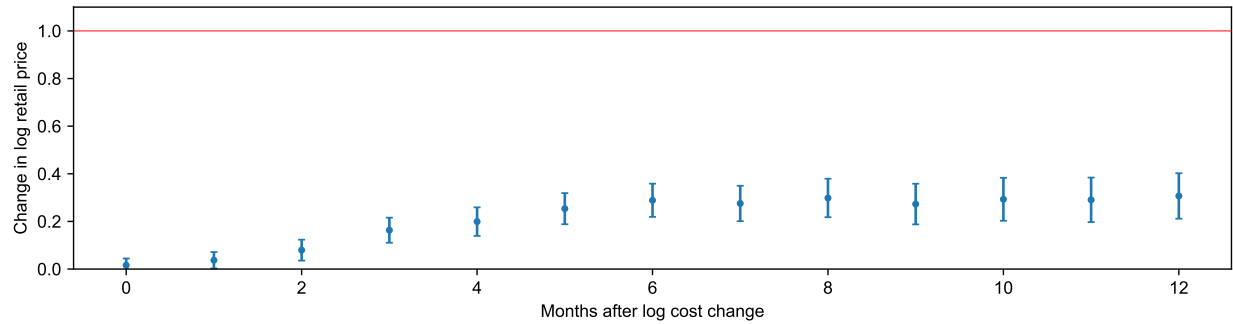
(c) Pass-through in logs.

Note: Panel (a) plots the time series of the commodity price from the IMF and the Average Price Data series from the BLS. The series are adjusted by the conversion factors in Appendix Table A4 so that the two series are in comparable units. Panels (b) and (c) plot the cumulative pass-through to month T, $\sum_{k=0}^{T} b_k$, from the specifications (2) and (3), using a total horizon of $K = 12$ months.

Figure A10: Passthrough of flour commodity costs to retail prices.
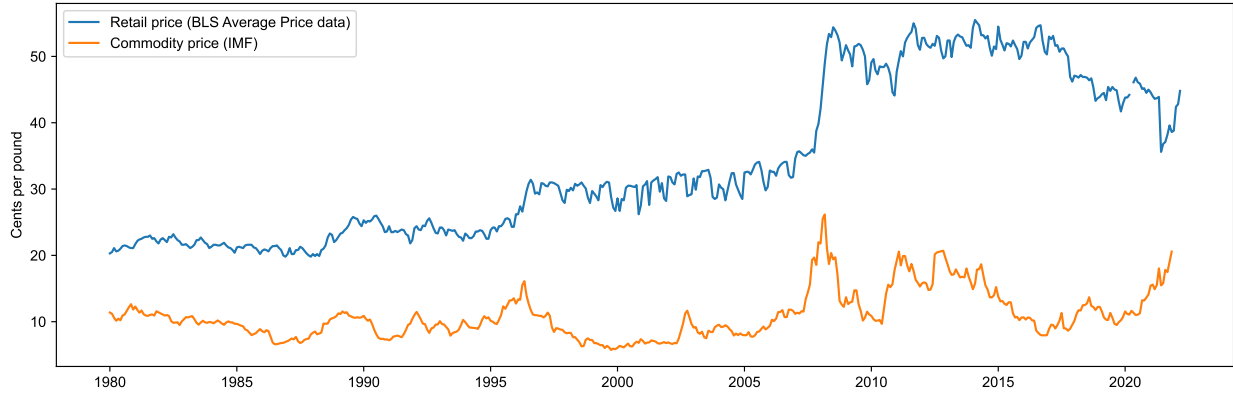
(a) Wheat commodity costs (IMF) and retail all-purpose flour prices (U.S. CPI).
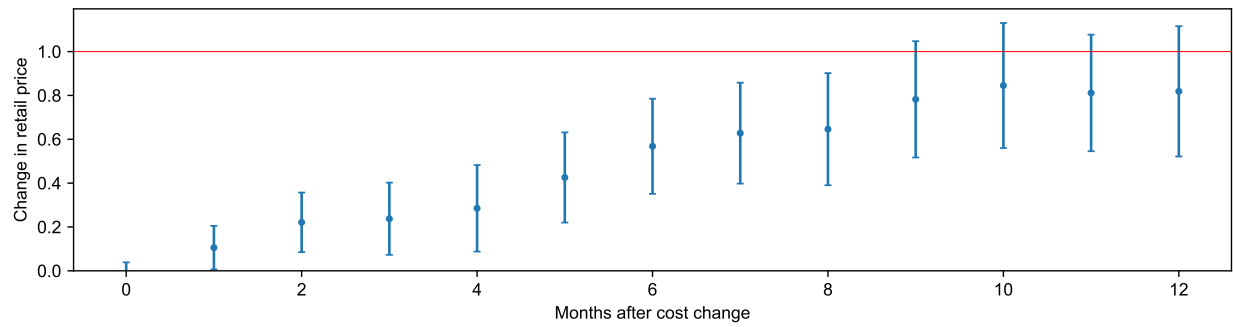
(b) Pass-through in levels.


(c) Pass-through in logs.

Note: Panel (a) plots the time series of the commodity price from the IMF and the Average Price Data series from the BLS. The series are adjusted by the conversion factors in Appendix Table A4 so that the two series are in comparable units. Panels (b) and (c) plot the cumulative pass-through to month T, $\sum_{k=0}^{T} b_k$, from the specifications (2) and (3), using a total horizon of $K = 12$ months.

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Figure A11: Passthrough of frozen orange juice commodity costs to retail prices.

(a) Frozen orange juice commodity costs (IMF) and retail orange concentrate prices (U.S. CPI).

(b) Pass-through in levels.
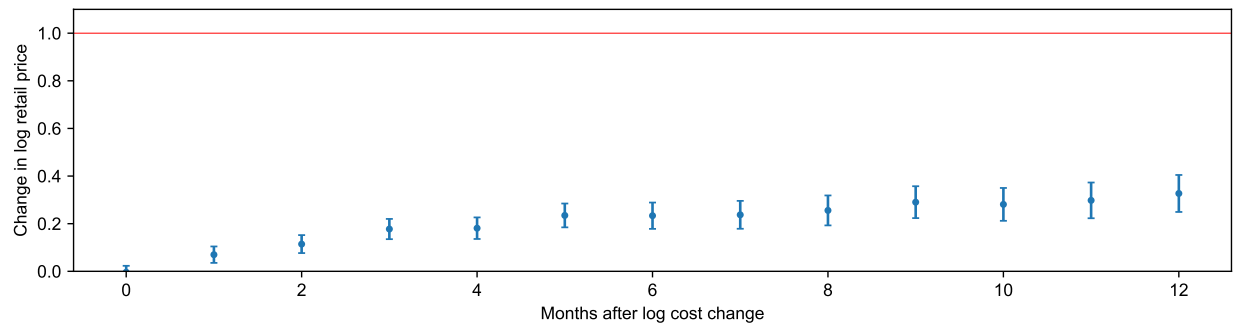
(c) Pass-through in logs.

Note: Panel (a) plots the time series of the commodity price from the IMF and the Average Price Data series from the BLS. The series are adjusted by the conversion factors in Appendix Table A4 so that the two series are in comparable units. Panels (b) and (c) plot the cumulative pass-through to month T, $\sum_{k=0}^{T} b_k$, from the specifications (2) and (3), using a total horizon of $K = 12$ months.

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Figure A12: Wholesale gas price and number of gasoline stations in Perth.
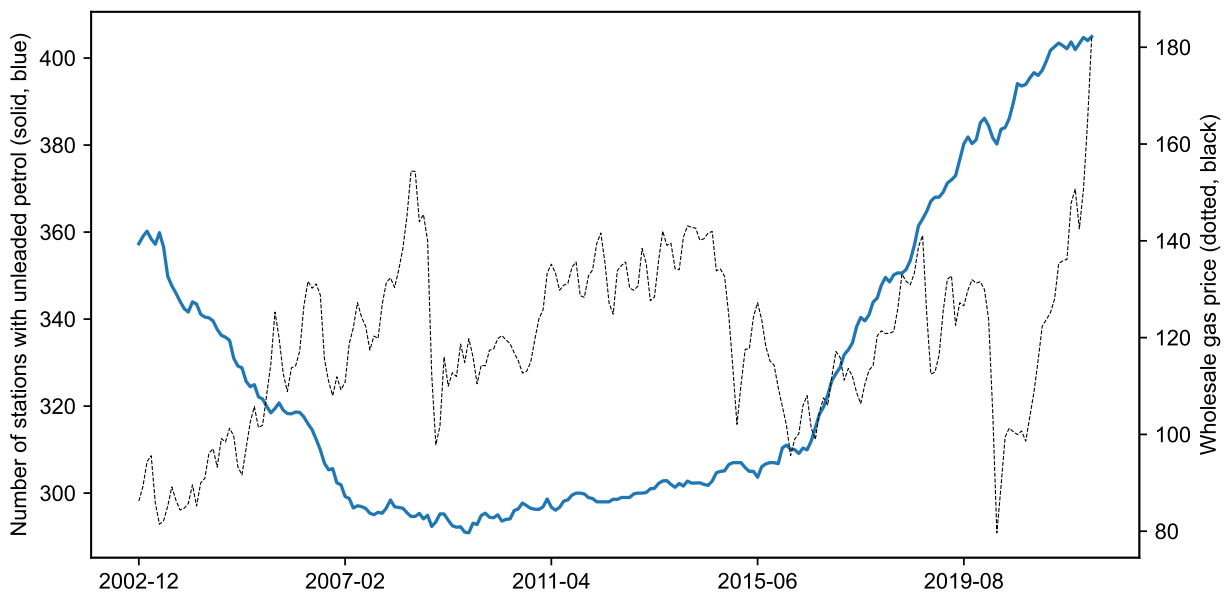
Note: Wholesale prices are TGP prices described in the text. The number of gas stations is the count of gas stations in the Perth metropolitan area with a non-missing unleaded petrol gas price.
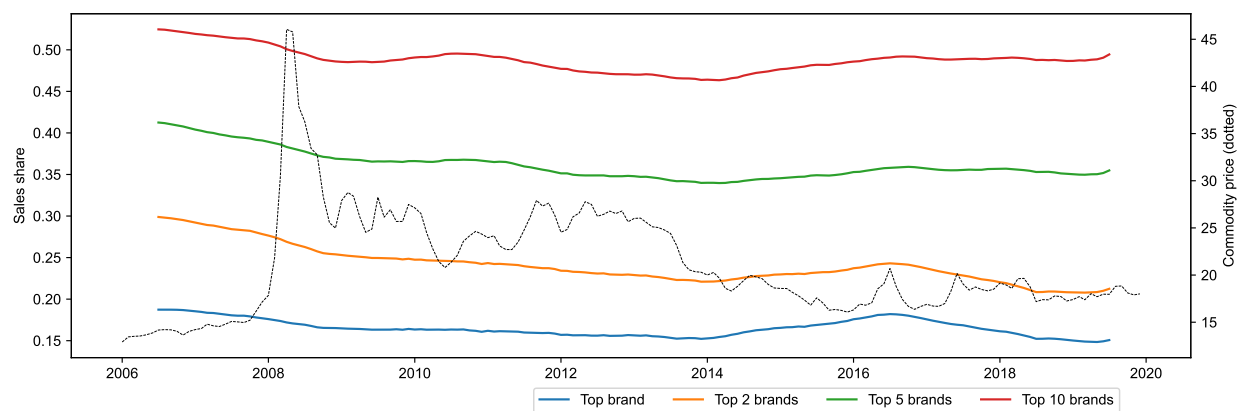
Figure A13: Commodity prices and market shares of top brands.
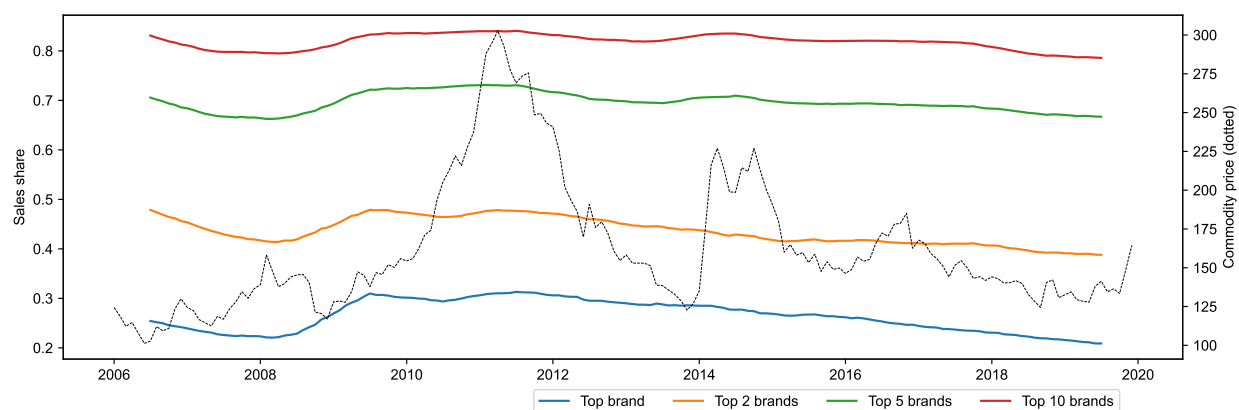
(a) Coffee
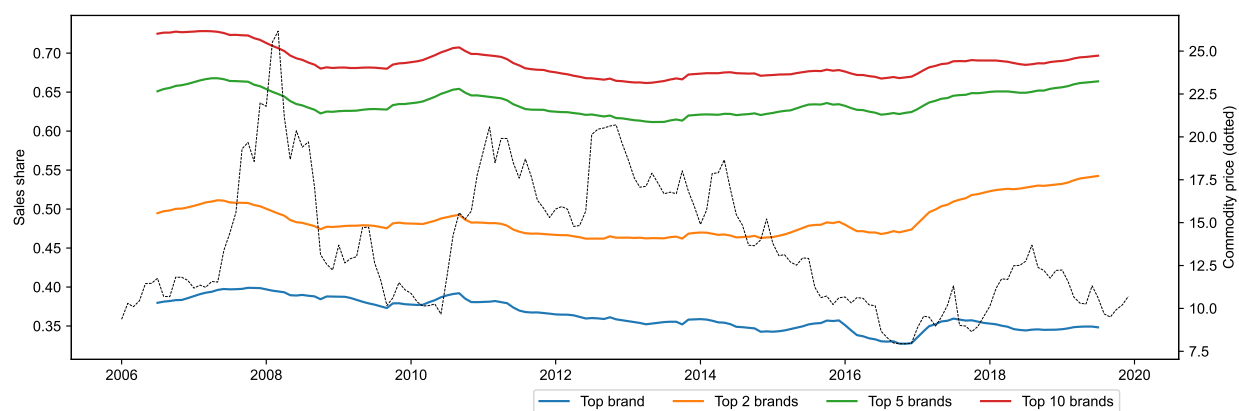
(b) Flour

(c) Rice

Note: Commodity prices are from the IMF. Brands are defined using unique brand identifiers provided by Nielsen. In each product module, brands are ranked by total sales over the full sample, and the share of sales by the top one, two, five, and ten brands is calculated as a six-month moving average of brand sales over total product module sales.

Figure A14: Food-at-home CPI, Food Manufacturing PPI, and Farm Products PPI.
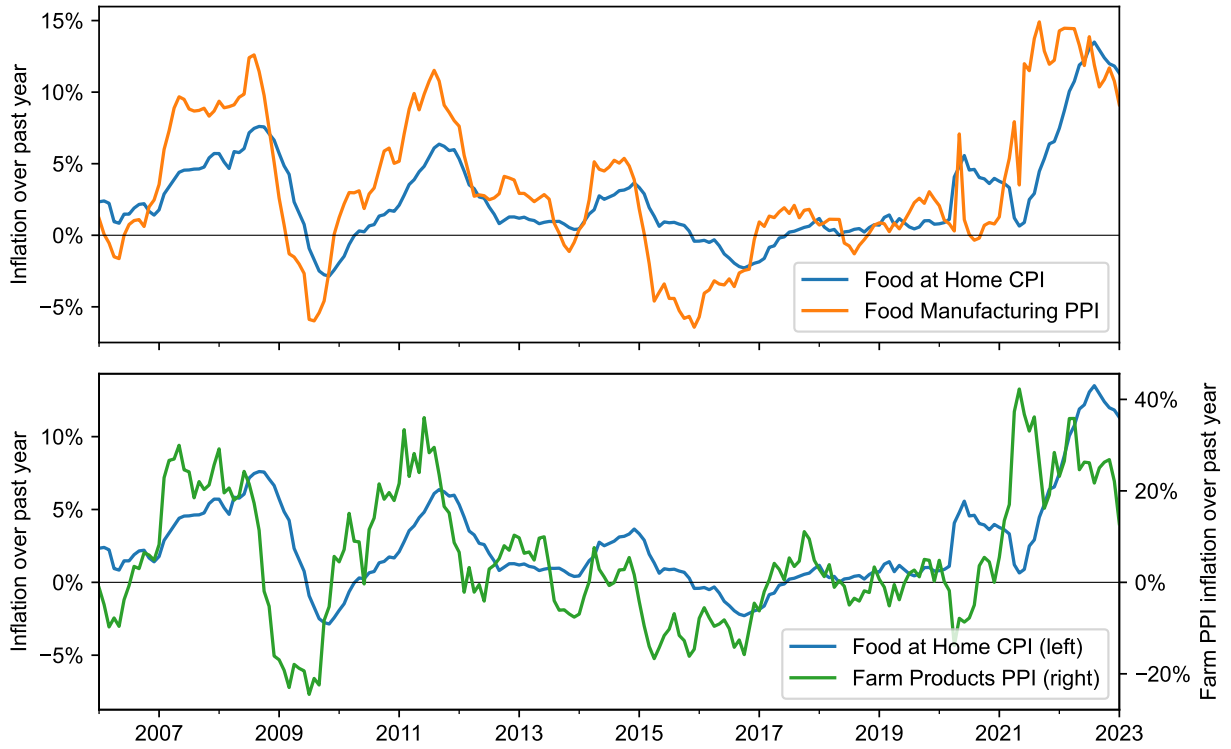
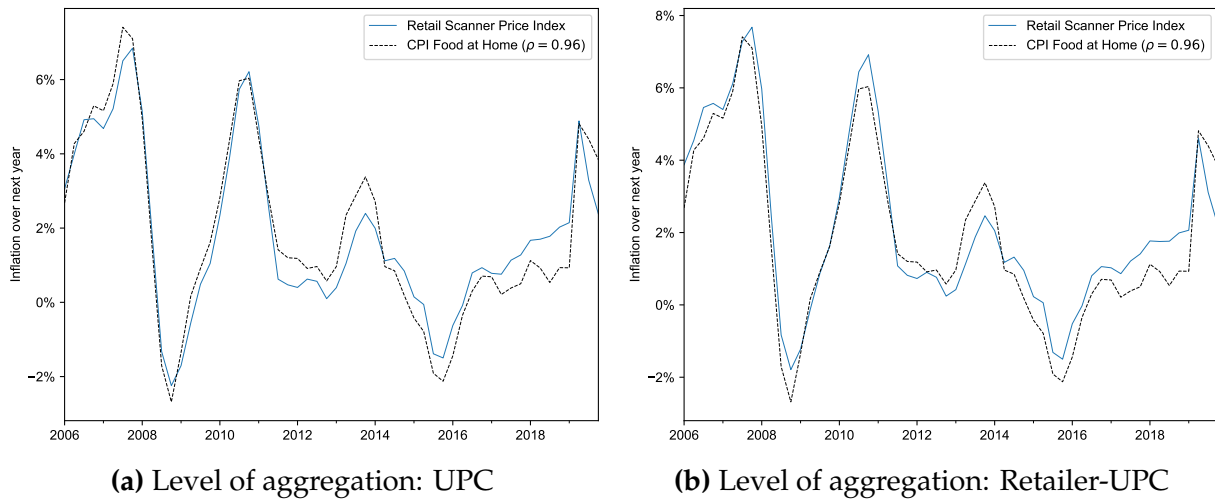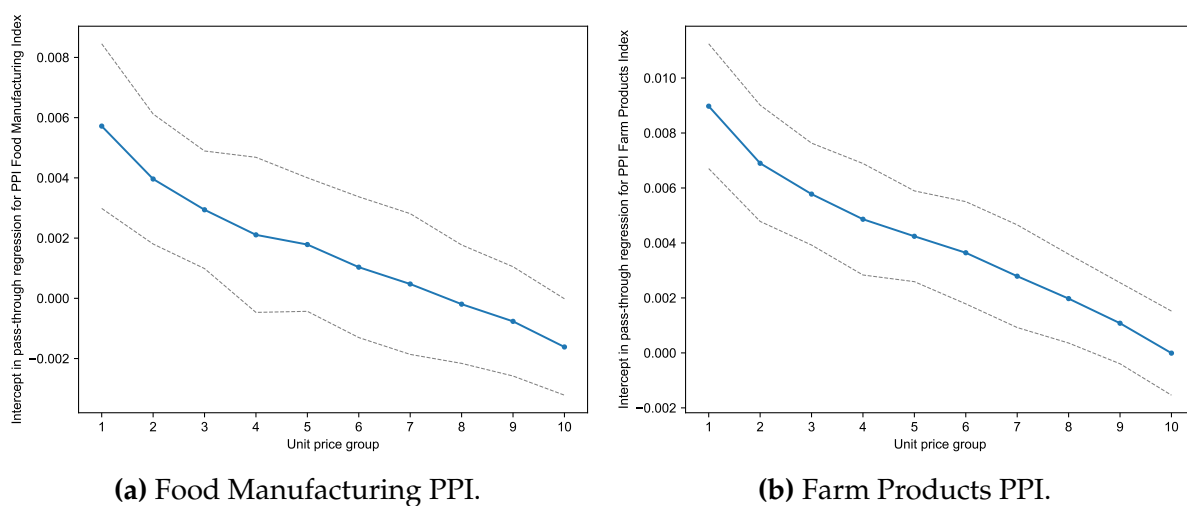**Figure A15:** Inflation rates on food at home CPI and Retail Scanner price index.

(a) Level of aggregation: UPC

(b) Level of aggregation: Retailer-UPC

Figure A16: Intercept in log pass-through regressions of upstream producer price indices.

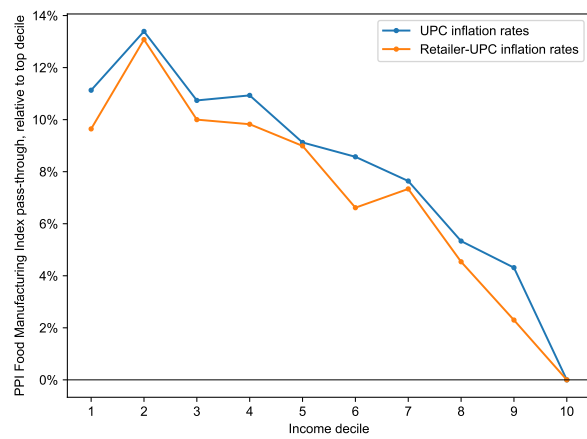**(a)** Food Manufacturing PPI.

(b) Farm Products PPI.

Note: Dotted lines indicate 95 percent confidence intervals using Driscoll-Kraay standard errors.

Figure A17: Predicted Retail Scanner index inflation using upstream PPIs and unemp. gap.
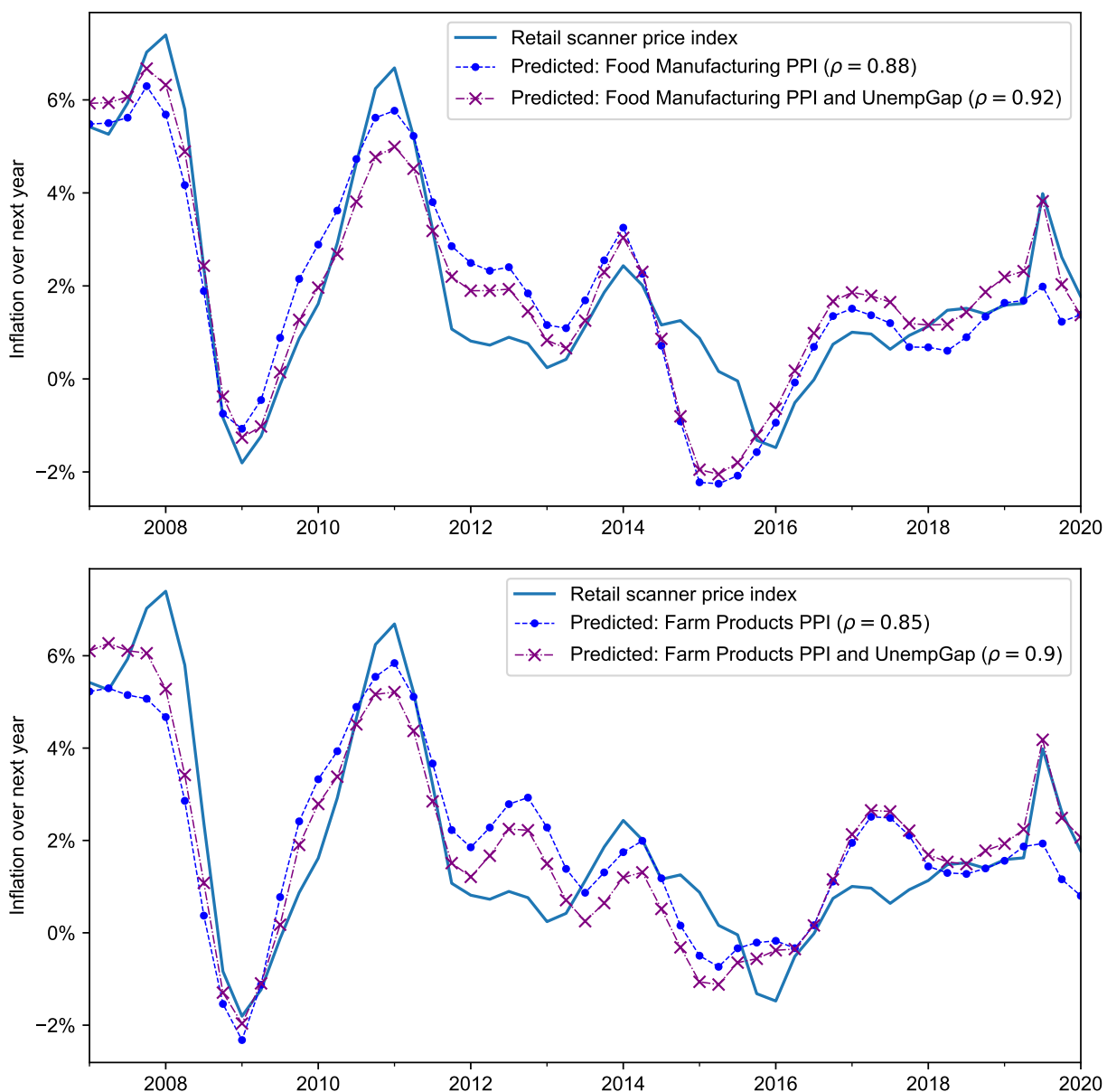

Note: The retail scanner index inflation plotted in each panel is constructed using retailer-UPC prices as the lowest level of aggregation. The top panel plots predicted values for year-over-year inflation from a long-run pass-through specification regressing changes in the retail scanner price index on four lags of Food Manufacturing PPI inflation and four lags of the unemployment gap. The coefficient ρ in the legend reports the correlation coefficient between the predicted values of year-over-year inflation and the actual retail scanner index inflation. The bottom panel repeats the exercise instead using Farm Products PPI inflation.

Figure A18: Differences by income decile: Sensitivity of food-at-home inflation to upstream PPI and variance of food-at-home inflation rates.

(a) Pass-through of Food Manufacturing PPI.
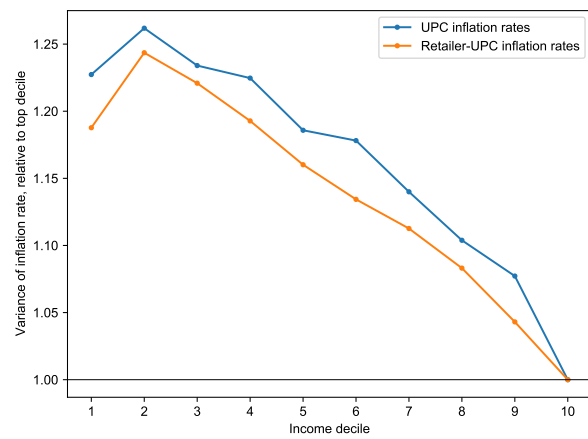
(b) Variance of inflation rates.

Appendix B   Proofs

B.1   Proofs for pass-through under relaxed assumptions

Let us consider how prices change (in levels) with a change in the commodity cost. Starting with a markup pricing rule,

$$p = \mu \times mc,$$

where mc is marginal cost, we totally differentiate with respect to the commodity cost c to get:

$$\frac{dp}{dc} = \frac{d\mu}{dp}\frac{dp}{dmc}\frac{dmc}{dc} \times mc + \mu \times \frac{dmc}{dc}$$

$$= \mu \left[\frac{d\log\mu}{d\log p}\frac{d\log p}{d\log mc} + 1 \right] \frac{dmc}{dc}.$$

Using the identity $d\log p = d\log\mu + d\log mc$, we simplify to get

$$\frac{dp}{dc} = \mu \left[\frac{1}{1 - \frac{d\log\mu}{d\log p}} \right] \frac{dmc}{dc}.$$

In my most flexible generalization, changes in marginal cost in response to changes in the commodity cost can come about from three channels: (1) direct effects, (2) correlation between changes in the commodity cost and costs of the other variable input, and (3) non-constant returns to scale. Writing this out:

$$\frac{dp}{dc} = \mu \left[\frac{1}{1 - \frac{d\log\mu}{d\log p}} \right] \left[\underbrace{\frac{\partial mc}{\partial c}}_{(1)} + \underbrace{\frac{\partial mc}{\partial w}\frac{dw}{dc}}_{(2)} + \underbrace{\frac{\partial mc}{\partial y}\frac{dy}{dp}\frac{dp}{dc}}_{(3)} \right].$$

Solving for the fixed point in dp/dc yields,

$$\frac{dp}{dc} = \frac{\mu}{1 - \frac{d\log\mu}{d\log p} + \sigma\frac{\partial\log mc}{\partial\log y}} \left(\frac{\partial mc}{\partial c} + \frac{\partial mc}{\partial w}\frac{dw}{dc} \right).$$

Let us assume that the markup is set using the Lerner pricing rule, $\mu = \sigma/(\sigma - 1)$. Then, we finally get

$$dp = \underbrace{\frac{\mu}{1 + \frac{1}{\sigma-1}\frac{d\log\sigma}{d\log p} + \sigma\frac{\partial\log mc}{\partial\log y}} \left(\frac{\partial mc}{\partial c} + \frac{\partial mc}{\partial w}\frac{dw}{dc}\right)}_{\text{Pass-through in levels}} dc.$$

The bracketed expression is the pass-through in levels, which is equal to one under complete pass-through in levels. We can see that relative to baseline case considered in the main text (constant returns to scale, constant elasticity of demand, uncorrelated costs, and Leontief production), a positive super-elasticity of demand or decreasing returns to scale each lead to a decline in pass-through in levels, while a positive correlation between the commodity cost and other non-commodity inputs results in an increase in pass-through in levels.

How marginal costs change with the commodity price, the non-commodity input's price, and scale requires writing down a production function for output. I consider the generalized production function,

$$y = \left(\omega x^{\frac{\theta-1}{\theta}} + (1-\omega)\ell^{\alpha\frac{\theta-1}{\theta}}\right)^{\frac{\theta}{\theta-1}},$$

where y is total output, x and ℓ are commodity and other variable inputs, θ is the elasticity of substitution between the commodity and the other variable input, ω are weights on the usage of the two inputs, and α determines returns to scale in use of the non-commodity input.

To consider how relaxing the assumptions in the main text affect pass-through in levels, I consider the following in turns: (1) a non-zero super-elasticity of demand $\frac{d\log\sigma}{d\log p} \neq 0$, (2) non-Leontief production ($\theta \neq 0$), (3) correlated costs ($\frac{dw}{dc} \neq 0$), and (4) decreasing returns to scale $\alpha < 1$.

Super-elasticity of demand. When the super-elasticity of demand is non-zero (but all other baseline assumptions hold), we get:

$$dp = \frac{\mu}{1 + \frac{1}{\sigma-1}\frac{d\log\sigma}{d\log p}} dc$$

$$= \frac{\sigma}{\sigma - 1 + \frac{d\log\sigma}{d\log p}} dc.$$

To get complete pass-through in levels, it is clear that we need $\frac{d\log\sigma}{d\log p} = 1$. This possibility is investigated empirically in the main text. A number of previous papers note that semilog demand curves ($\log y = -\alpha p + C$) yield complete pass-through in levels (Bulow and Pfleiderer 1983; Weyl and Fabinger 2013; Mrázová and Neary 2017). To see this, note that the elasticity of demand under semilog demand curves is

$$\sigma = -\frac{d\log y}{d\log p} = \alpha p,$$

and hence $\frac{d\log\sigma}{d\log p} = 1$.

Non-Leontief production. Suppose $\theta \neq 0$. Then, we find

$$dp = \mu\left(\frac{\partial mc}{\partial c}\right)dc = \mu\left(\frac{c}{\omega C}\right)^{-\theta}dc.$$

For complete pass-through in levels, we must have

$$\theta = \frac{\log\mu}{\log\frac{c}{\omega C}},$$

which cannot hold always, since the ratio c/C changes with fluctuations in the commodity cost c.

Correlated costs. Suppose $\frac{dw}{dc} \neq 0$. Then, we find

$$dp = \mu\left(1 + \frac{dw}{dc}\right)dc.$$

Complete pass-through in levels requires

$$\frac{dw}{dc} = \frac{1-\mu}{\mu} = -\frac{1}{\sigma}.$$

This is unlikely to explain complete pass-through in levels across products or markets, since elasticities of demand for products within a market tend to exhibit considerable variation. Moreover, in most environments we expect input costs to be positively correlated. For example, in the market for retail gasoline, other variable inputs like shipping/transportation costs are likely to be increasing in the cost of gas.

Decreasing returns to scale. Suppose $\alpha < 1$. Pass-through in levels is then

$$dp = \frac{\mu}{1 + \sigma \frac{\partial \log mc}{\partial \log y}} dc,$$

which means that complete pass-through in levels requires

$$\frac{\partial \log mc}{\partial \log y} = \frac{\mu - 1}{\sigma} > 0.$$

In terms of production function primitives, the elasticity of marginal costs to output is

$$\frac{\partial \log mc}{\partial \log y} = \left(\frac{1 - \alpha}{\alpha}\right) \frac{w \frac{1}{\alpha} y^{\frac{1-\alpha}{\alpha}}}{c + w \frac{1}{\alpha} y^{\frac{1-\alpha}{\alpha}}}.$$

It is clear that we need $\alpha < 1$, or decreasing returns to scale, to generate $\frac{\partial \log mc}{\partial \log y} > 0$. Note however that the elasticity of marginal costs to output depends on the commodity cost c and the level of output y. It is not possible to choose α so that $\frac{\partial \log mc}{\partial \log y}$ is positive and constant as the commodity cost c fluctuates.

B.2 Proofs for safety margin model

The firm maximization problem is

$$\max_{p} \mathbb{E}\left[\pi(p)\right],$$

where profits are

$$\pi(p) = (p - c - w) D(p) - f_o N^{-\zeta}., \quad \text{and} \quad D(p_t) = \varepsilon \frac{1}{N}\left(\frac{p}{P}\right)^{-\sigma},$$

subject to the safety margin constraint

$$\Pr\left[\pi(p) \leq 0\right] \leq \phi, \quad \forall c, w, f_o.$$

Dixit Stiglitz equilibrium. Solving the first order condition yields the optimal price when the constraint does not bind,

$$p_t^{\text{DS}} = \frac{\sigma}{\sigma - 1}(c_t + w_t).$$

Plugging this into the expression for profits above, firms' expected profits are

$$\mathbb{E}\pi(p_t) = \frac{1}{\sigma - 1}\,(c_t + w_t)\,\frac{1}{N_t^{\text{DS}}} - w_t f_o \left(\frac{1}{N_t^{\text{DS}}}\right)^{\zeta}.$$

With free entry, discounted expected future profits equal the cost of entry $w_t f_e$. That is,

$$\mathbb{E}_t \sum_{k=0}^{\infty} \beta^k \pi_{t+k} = w_t f_e.$$

Plugging in the expression for profits, we get:

$$\mathbb{E}_t \sum_{k=0}^{\infty} \frac{w_{t+k}}{w_t} \beta^k \left[\frac{1}{\sigma - 1}\left(\frac{c_{t+k}}{w_{t+k}} + 1\right)\frac{1}{N_{t+k}^{\text{DS}}} - f_o\left(\frac{1}{N_{t+k}^{\text{DS}}}\right)^{\zeta}\right] = f_e.$$

Using the assumption that wages are expected to grow at a constant rate g and that the ratio of the commodity cost to the wage c_t/w_t is a random walk with zero drift, we can see that one solution to this equation is to set

$$\frac{1}{\sigma - 1}\left(\frac{c_t}{w_t} + 1\right)\frac{1}{N_t^{\text{DS}}} - f_o\left(\frac{1}{N_t^{\text{DS}}}\right)^{\zeta} = f_e\left[1 - \beta\left(1 + g\right)\right].$$

Solving, we find that the number of firms satisfies

$$N_t^{\text{DS}} f_e \left[1 - \beta\left(1 + g\right)\right] + f_o \left(N_t^{\text{DS}}\right)^{1-\zeta} = \frac{1}{\sigma - 1}\left(1 + \frac{c_t}{w_t}\right).$$

By taking the derivative with respect to c_t/w_t, we can see that the number of firms N_t^{DS} increases when the commodity price is high relative to the non-commodity input price:

$$\frac{dN_t^{\text{DS}}}{d\left(\frac{c_t}{w_t}\right)} = \frac{1}{\sigma - 1}\left[f_e\left[1 - \beta\left(1 + g\right)\right] + f_o\left(N_t^{\text{DS}}\right)^{-\zeta}\right]^{-1} > 0.$$

The gross margin m_t is equal to total sales minus variable costs of goods sold as a fraction of sales, which is

$$m_t^{\text{DS}} = \frac{p_t^{\text{DS}} y_t - (c_t + w_t) y_t}{p_t^{\text{DS}} y_t} = \frac{1}{\sigma}.$$

Equilibrium when the safety constraint binds.   When the safety constraint is binding, we have:

$$\Pr\left[\pi(p_t^{\text{safe}}) \leq 0\right] = \phi.$$

Substituting in for profits and rearranging, we get the condition:

$$\Pr\left[\varepsilon \leq \frac{w_t f_o \left(N_t^{\text{safe}}\right)^{1-\zeta}}{p_t^{\text{safe}} - c_t - w_t}\right] = \phi.$$

Let H be the CDF of the demand shock ε, and H^{-1} be its inverse. We get the pricing rule

$$p_t^{\text{safe}} = c_t + w_t + \frac{w_t f_o \left(N_t^{\text{safe}}\right)^{1-\zeta}}{H^{-1}\left(\phi\right)}.$$

As before, we use the free entry condition to solve for N_t^{safe}:

$$\mathbb{E}_t \sum_{k=0}^{\infty} \beta^k \left[(p_{t+k} - c_{t+k} - w_{t+k}) \frac{1}{N_{t+k}^{\text{safe}}} - w_{t+k} f_o \left(N_{t+k}^{\text{safe}}\right)^{-\zeta}\right] = w_t f_e.$$

Solving yields,

$$f_o \left[\frac{1}{H^{-1}\left(\phi\right)} - 1\right] \sum_{k=0}^{\infty} \beta^k \left(1 + g\right)^k \left(N_{t+k}^{\text{safe}}\right)^{-\zeta} = f_e.$$

A solution to this equation is $N_{t+k}^{\text{safe}} = N^{\text{safe}}$, where

$$\left(N^{\text{safe}}\right)^{-\zeta} = \frac{\Delta f_e}{f_o} \frac{H^{-1}\left(\phi\right)}{1 - H^{-1}\left(\phi\right)}.$$

Note here that it becomes important to have $\zeta > 0$, since at $\zeta = 0$ it is impossible to satisfy this condition. Solving for the number of firms, we get the constant:

$$N^{\text{safe}} = \left[\frac{f_o}{\Delta f_e} \frac{1 - H^{-1}\left(\phi\right)}{H^{-1}\left(\phi\right)}\right]^{\frac{1}{\zeta}}.$$

Using our expression for the price p_t^{safe}, we find that gross margins decrease in c_t/w_t:

$$m_t^{\text{safe}} = \frac{1}{1 + \left(1 + \frac{c_t}{w_t}\right)\left[\frac{1 - H^{-1}\left(\phi\right)}{\Delta f_e N^{\text{safe}}}\right]}.$$

When does the safety constraint bind?   The safety margin constraint binds when the Dixit-Stiglitz pricing rule violates the safety margin constraint.

$$\Pr\left[\pi(p_t^{\text{DS}}) \le 0\right] \ge \phi.$$

Plugging in our expression for profits in the Dixit-Stiglitz equilibrium and simplifying, we find that the constraint is binding if

$$\Pr\left[\varepsilon \le \frac{f_o}{f_o + \Delta f_e \left(N_t^{\text{DS}}\right)^\zeta}\right] \ge \phi.$$

In the case where $\zeta = 0$, this is simply,

$$\Pr\left[\varepsilon \le \frac{f_o}{f_o + \Delta f_e}\right] \ge \phi.$$

When $\zeta > 0$, the likelihood that the constraint binds is decreasing in N_t^{DS}, and since N_t^{DS} is increasing in c_t/w_t, is decreasing in c_t/w_t. This monotonicity with respect to c_t/w_t means that there is a single cutoff c^* such that the constraint only binds if $c_t/w_t \le c^*$.

At the cutoff c^*, the safety margin constraint is exactly binding. We will now explore the comparative static of c^* with respect to exogenous parameters of the model. Using H to denote the CDF of ε, the cutoff c^* satisfies

$$H\left(\frac{f_o}{f_o + \Delta f_e \left(N^{\text{DS}}\left(c^*\right)\right)^\zeta}; Var(\varepsilon)\right) = \phi,$$

where $Var(\varepsilon)$ is the variance of ε and indexes the function H. An increase in $Var(\varepsilon)$ increases the mass to the left of any given point, so $\partial H/\partial Var(\varepsilon) > 0$. First, taking the comparative static of c^* with respect to $Var(\varepsilon)$, we find

$$\frac{dc^*}{dVar(\varepsilon)} = \frac{-\frac{\partial H}{\partial Var(\varepsilon)}}{H'\left(x\right)\frac{dx}{dc^*}} > 0.$$

where we use the shorthand $x \equiv \frac{f_o}{f_o + \Delta f_e \left(N^{\text{DS}}(c^*)\right)^\zeta}$ and note that $dx/dc^* < 0$. Intuitively, increasing the variance of the demand shock increases the likelihood that the firms' variable profits will fall short of costs, and thus increases the likelihood of the safety constraint

binding. Similarly, taking the derivative with respect to ϕ yields,

$$\frac{dc^*}{d\phi} = \frac{1}{H'(x)\frac{dx}{dc^*}} < 0.$$

Intuitively, a manager who is more tolerant of the risk of negative profits (as ϕ increases) is less likely to follow safety pricing.

Taking the comparative static with respect to f_o and Δf_e requires differentiating both the cutoff condition and the condition that pins down the number of firms N_t^{DS}. First, for f_o, differentiating the condition for the number of firms yields,

$$dN_t^{DS}\left[\Delta f_e + f_o(1 - \zeta)\left(N_t^{DS}\right)^{-\zeta}\right] = \frac{1}{\sigma - 1}dc^* - df_o\left(N_t^{DS}\right)^{1-\zeta}.$$

Differentiating the cutoff condition yields

$$H'(x)\left[df_o - \frac{f_o}{f_o + \Delta f_e\left(N^{DS}(c^*)\right)^{\zeta}}\left[df_o + \Delta f_e\zeta\left(N^{DS}\right)^{\zeta-1}dN^{DS}\right]\right] = 0.$$

Combining the two conditions and simplifying yields

$$\left[\frac{\Delta f_e\left(N^{DS}(c^*)\right)^{\zeta}}{f_o} + \frac{\zeta\Delta f_e}{\Delta f_e + f_o(1 - \zeta)\left(N_t^{DS}\right)^{-\zeta}}\right]df_o = \frac{\Delta f_e\zeta\left(N^{DS}\right)^{\zeta-1}}{\Delta f_e + f_o(1 - \zeta)\left(N_t^{DS}\right)^{-\zeta}}\left[\frac{1}{\sigma - 1}dc^*\right].$$

The coefficients in front of both df_o and dc^* are positive, and hence c^* is increasing in f_o. Intuitively, when overhead costs are high, there is a greater chance that variable profits will not be able to cover the fixed costs of operation, and hence the safety constraint is more likely to bind.

Finally, for Δf_e. Differentiating the condition for the number of firms yields,

$$\left[\Delta f_e + (1 - \zeta)f_o\left(N_t^{DS}\right)^{-\zeta}\right]dN_t^{DS} = -N_t^{DS}d(\Delta f_e) + \frac{1}{\sigma - 1}dc^*.$$

Differentiating the cutoff condition yields

$$d(\Delta f_e)\left(N^{DS}\right)^{\zeta} + \zeta\Delta f_e\left(N^{DS}\right)^{\zeta-1}dN^{DS} = 0.$$

Combining the two and simplifying, we get

$$\frac{1-\zeta}{\zeta}\frac{\Delta f_e N^{\text{DS}} + f_o\left(N_t^{\text{DS}}\right)^{1-\zeta}}{\Delta f_e} d\left(\Delta f_e\right) = -\frac{1}{\sigma-1}dc^*.$$

Thus, for $\zeta \in (0,1)$, $dc^*/df_e < 0$.

Comparing prices and number of firms in the two equilibria. When the constraint binds, how do the safety margin equilibrium prices and number of firms compare to the Dixit-Stiglitz benchmarks? Recall that when the safety margin constraint binds, we have

$$\Pr\left[\varepsilon \leq \frac{f_o}{f_o + \Delta f_e\left(N_t^{\text{DS}}\right)^{\zeta}}\right] \geq \phi.$$

This means

$$H^{-1}\left(\phi\right) \leq \frac{f_o}{f_o + \Delta f_e\left(N_t^{\text{DS}}\right)^{\zeta}}.$$

Rearranging, we get

$$\left(N_t^{\text{DS}}\right)^{\zeta} \leq \frac{f_o}{\Delta f_e}\frac{1 - H^{-1}\left(\phi\right)}{H^{-1}\left(\phi\right)} = \left(N^{\text{safe}}\right)^{\zeta}.$$

Hence, $N_t^{\text{DS}} \leq N^{\text{safe}}$. To compare prices in both equilibria, we take the difference

$$p_t^{\text{safe}} - p_t^{\text{DS}} = w_t\left(\frac{f_o\left(N^{\text{safe}}\right)^{1-\zeta}}{H^{-1}\left(\phi\right)} - \frac{1}{\sigma-1}\left(\frac{c_t}{w_t}+1\right)\right)$$

$$\geq w_t\frac{f_o}{H^{-1}\left(\phi\right)}\left(\left(N^{\text{safe}}\right)^{1-\zeta} - \left(N_t^{\text{DS}}\right)^{1-\zeta}\right) > 0.$$

where in the last line, we used the fact that when the safety constraint binds,

$$\Pr\left[\varepsilon \leq 1 - \frac{(\sigma-1)\Delta f_e}{1 + \frac{c_t}{w_t}}N_t^{\text{DS}}\right] \geq \phi,$$

and hence,

$$\frac{1 + \frac{c_t}{w_t}}{\sigma-1} \leq \frac{f_o\left(N_t^{\text{DS}}\right)^{1-\zeta}}{H^{-1}\left(\phi\right)}.$$

Thus, we conclude $p_t^{\text{safe}} \geq p_t^{\text{DS}}$.

Appendix C Retail Gasoline Evidence from Other Markets

C.1 Canada

I use weekly price data for 71 cities in 10 Canadian provinces provided by Kalibrate solutions.[51] These prices are collected across cities through a daily survey of pump prices funded by the Government of Canada and used for analyses by National Resources Canada. As an example, Figure C1 shows crude oil prices, wholesale prices, and retail prices excluding taxes for the City of Toronto.

Figure C1: Retail (excl. taxes), wholesale, and crude prices for the City of Toronto.
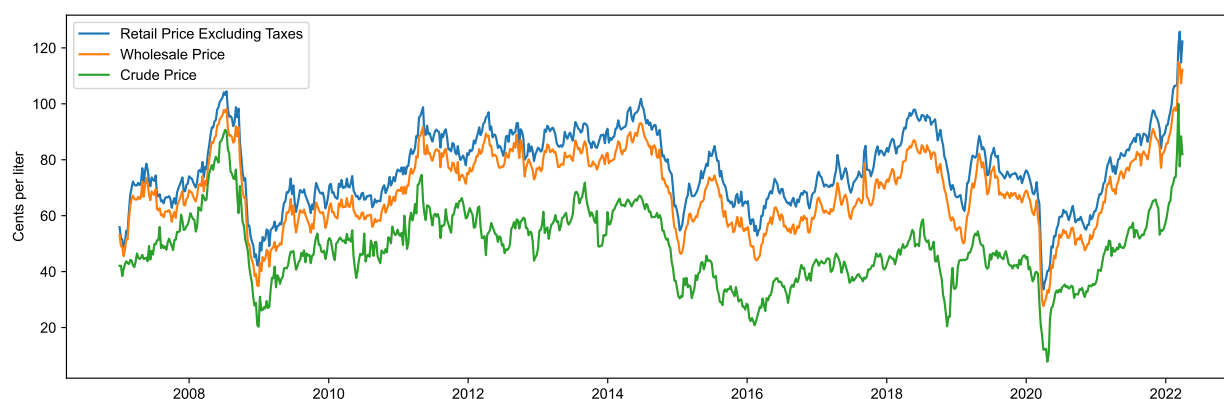

Figure C2 and Figure C3 plot pass-through from zero to eight weeks estimated in both logs and levels; the long-run pass-through estimates are reported in the main text in Table 4. For both parts of the supply chain, pass-through from costs to prices is complete in levels, though it is incomplete in logs.

C.2 South Korea

I download daily station-level price data from Opinet, a service started in 2008 by the Korea National Oil Corporation to provide customer transparency about petroleum product prices and enable research.[52] These data cover all gas stations within each city in South Korea; data files are available by city/county within each province. However, some stations appear to have incomplete coverage. Hence, for all results using these data, I limit my analyses to stations that have at least 500 daily price observations (i.e., at least 10% of days during the full sample period). Opinet also provides weekly average refinery

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[51] Weekly prices can be downloaded from `https://charting.kalibrate.com`.

[52] These data are available for download at `https://www.opinet.co.kr`.

Figure C2: Passthrough of Canadian crude prices to wholesale prices: Levels (top) and logs (bottom).
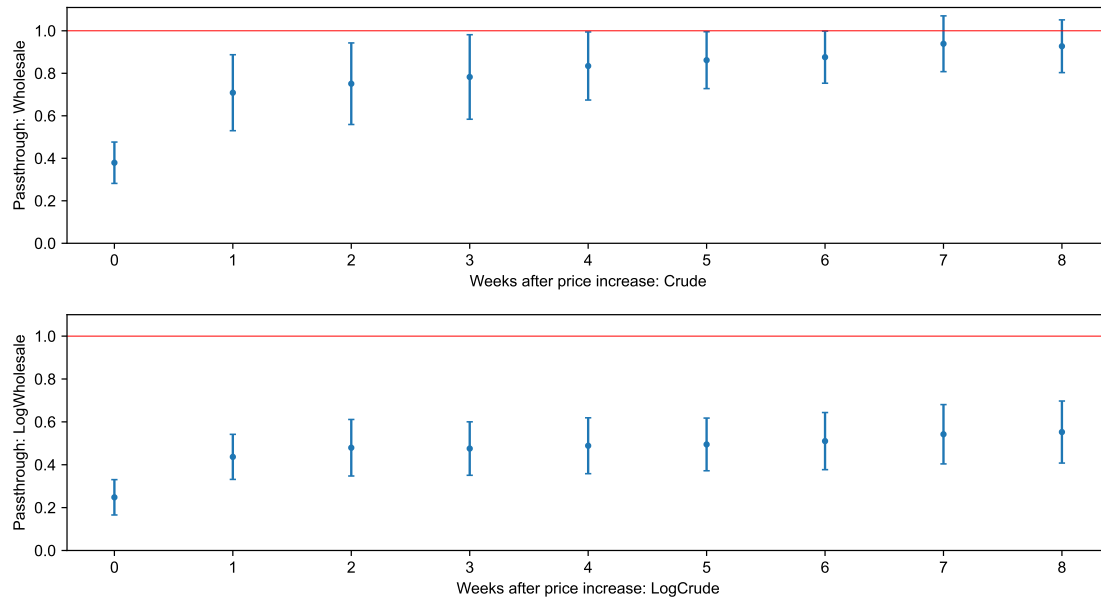

Figure C3: Passthrough of Canadian wholesale prices to retail prices excluding taxes: Levels (top) and logs (bottom).
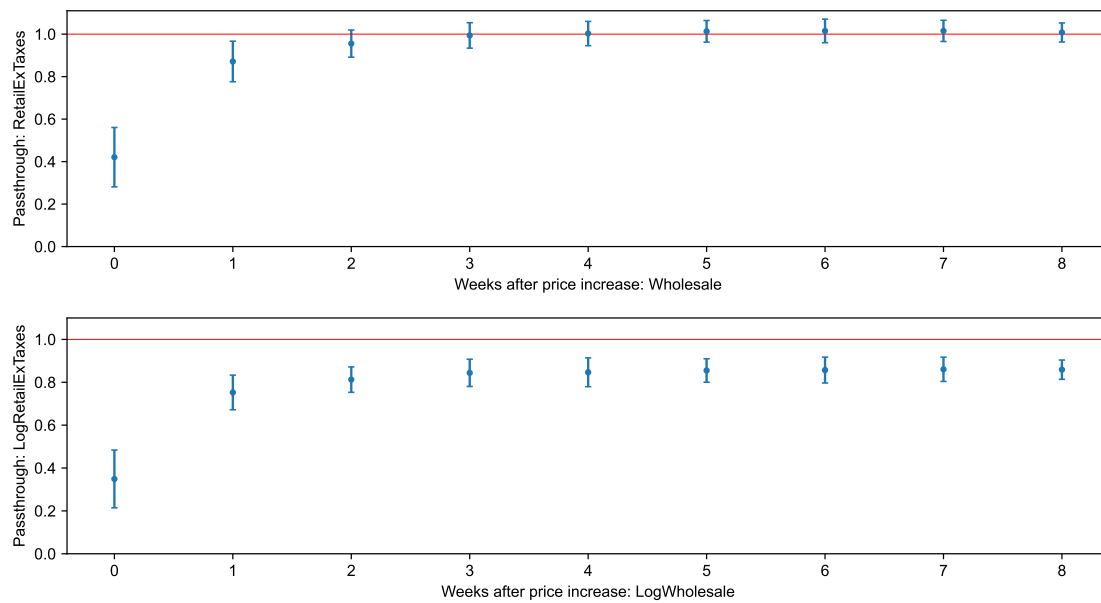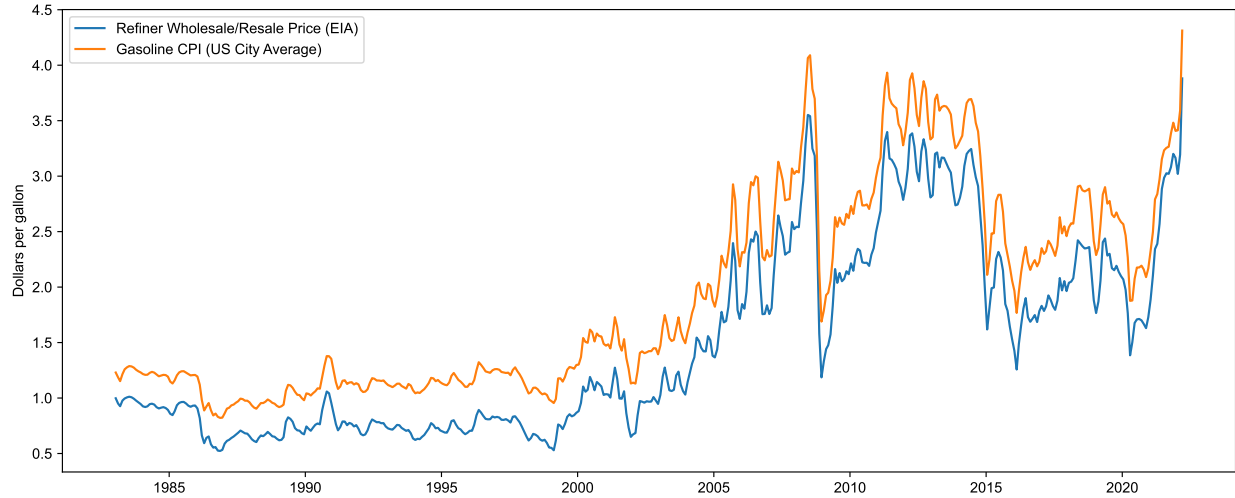
Figure C4: U.S. refiner wholesale/resale prices (EIA) and retail prices (BLS).

Note: Refiner wholesale prices are U.S. refiner gasoline prices for sale through retail outlets from the Energy Information Administration (EIA). Retail gasoline prices are from the BLS Average Price Data for the U.S. city average (Series ID APU000074714).

supply prices, which I use as the measure of costs facing retail stations. Estimates of the long-run pass-through from this data are reported in the main text in Table 4.

C.3   United States

National U.S. refinery and retail gasoline prices used in Figure A6 come from two sources: refinery prices are U.S. refiner gasoline prices for sale through retail outlets from the Energy Information Administration (EIA sourcekey EMA_EPMR_PTR_NUS_DPG), and retail prices are from the BLS Average Price Data for the U.S. city average (Series ID APU000074714). Figure C4 plots both time series.

Figure A6 calculates pass-through as the relationship between one-month price changes in refinery and retail prices. Estimates of pass-through at alternative horizons (i.e., 2-, 3-, 6-, and 12-month changes in refinery and retail prices) are nearly identical.

Appendix D Demand Elasticity Estimates

In each product category, I estimate demand elasticities separately for each UPC i at each store s. My baseline specification, which follows closely from DellaVigna and Gentzkow (2019), estimates the response of weekly log quantities to weekly log price, allowing for store-product-year and store-product-week-of-year fixed effects:

$$\log q_{i,s,t} = \eta \log p_{i,s,t} + \kappa (\log p_{i,s,t})^2 + \gamma \log p_{s,t}^{\text{store}} + \delta_{i,s,\text{year}(t)} + \alpha_{i,s,\text{week-of-year}(t)} + \varepsilon_{i,s,t}. \quad (15)$$

Following DellaVigna and Gentzkow (2019), I address the endogeneity of prices by instrumenting for price of UPC i at store s using the price of i at stores in the same retail chain as s, but outside s's geographic market (DMA).[53] These Hausman (1996) instruments are strongly correlated with true prices, due to retailers' tendencies to set uniform prices across locations, and hence have a strong first stage. Under the assumption that chain-level variation in prices are unrelated to local demand shocks at a specific store, it also avoids the endogeneity that would attenuate our estimates of the demand elasticity η.

There are two differences between (15) and the estimating equation in DellaVigna and Gentzkow (2019). First, (15) adds the squared log price, $(\log p_{i,s,t})^2$, as an additional independent variable. This follows Burya and Mishra (2023), who show that the super-elasticity of demand is identified by κ/η. Note that I use the Hausman (1996) instrument described above and its square to instrument for both log prices and squared log prices when estimating (15). Second, I also add the log average unit price in i's product category, $\log p_{s,t}^{\text{store}}$. This addition is made to reflect the fact that many models of log-concave demand curves consider how the elasticity of demand for a product changes as the product's price deviates from the average across other products. Nevertheless, I find in practice that results from estimating (15) are very similar whether or not we control for the average store price.

Appendix E Evidence from Other Food Products

In this appendix, I document complete pass-through in levels in two additional datasets on food product markets. First, I explore pass-through in the beef and pork markets using aggregate price data from the USDA. The advantage of these data is that they document prices at the farm, wholesale, and retail levels, allowing me to explore how pass-through

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[53]DMAs (designated market areas) are large, non-overlapping geographic regions defined by Nielsen that typically include several counties.

behavior varies between farm to wholesale prices and wholesale to retail prices. In both parts of the market, I find evidence of complete pass-through in levels.

Second, I explore pass-through of corn farm prices to the prices of downstream products, such as corn starch, high-fructose corn syrup, and dextrose. I find that products that have a lower cost share of corn exhibit slower pass-through, but appear to converge to complete pass-through in levels at long horizons.

E.1   Beef and pork, from farm to retail

I use monthly data compiled by the US Department of Agriculture (USDA) on average farm, wholesale, and retail prices for beef and pork.[54]   The USDA collects these retail prices from the Bureau of Labor Statistics and farm and wholesale prices from Agricultural Marketing Service reports. Figure E1 plots the time series of beef and pork farm, wholesale, and retail prices. For both beef and pork, there is a growing gap over time between wholesale and retail prices. The USDA also includes an additional price series for fresh retail beef products.

Table E1 reports the long-run pass-through (at a horizon of 12 months) of upstream beef and pork prices to downstream prices at various links in the chain from farm price to retail price. Across nearly all links, the estimated pass-through in levels is statistically indistinguishable from one, while the estimated log pass-through is always significantly less than one.

E.2   Corn downstream products

Monthly prices for corn and downstream products come from the USDA's Feed Grains Outlook. Table E2 lists the time series available from the USDA: in addition to the farm price of corn, the USDA reports the price of corn at nine corn markets across the US and monthly prices for six downstream products.

When considering the pass-through of corn farm prices to downstream products, it is important to take into account that manufacturing corn syrup and other downstream products from corn through the wet milling process produces byproducts—corn gluten feed and corn gluten meal—that are sold as feed for livestock. Hence, the commodity cost of corn needs to be adjusted downward to take into account the sale of these byproducts.

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[54]These data are available for download at `https://www.ers.usda.gov/data-products/meat-price-spreads`. The USDA also tracks a broiler prices, but these are a composite price index that includes several types of poultry.

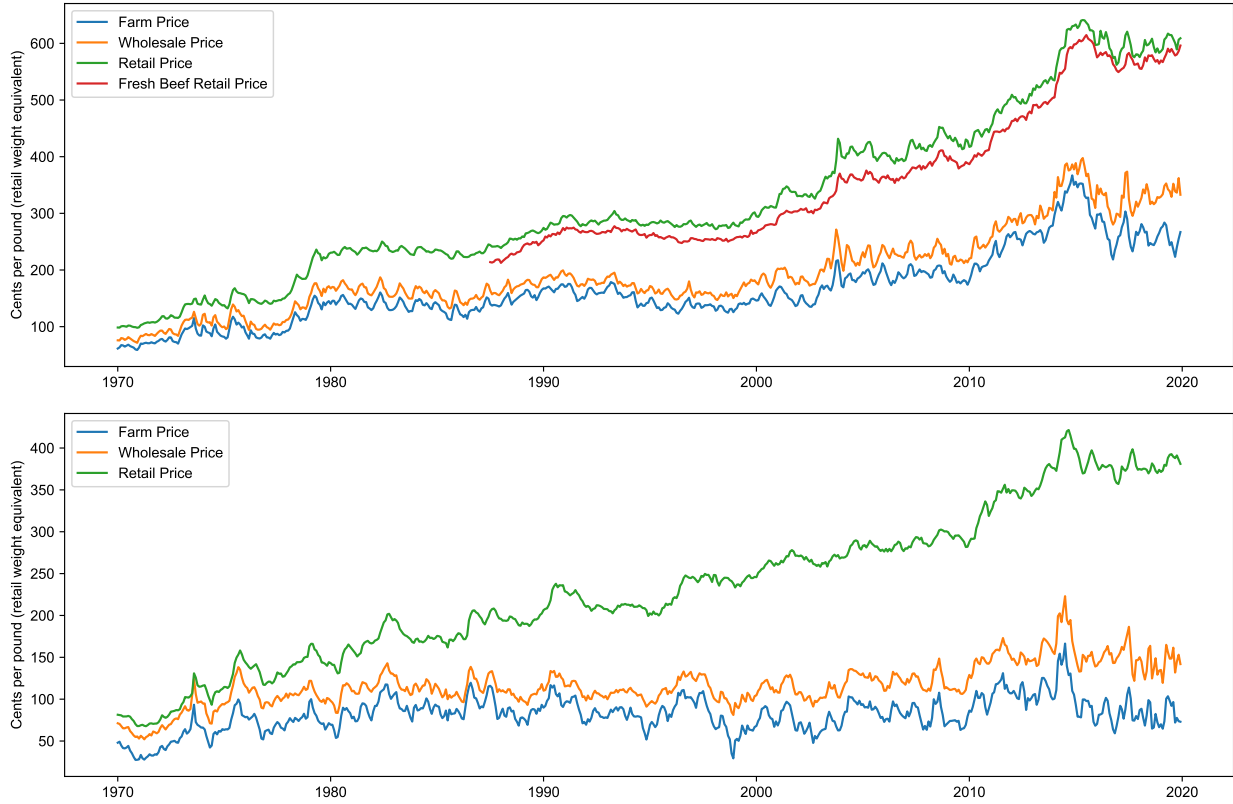**Figure E1:** Beef (top) and pork (bottom) prices over time.

Table E1: Pass-through of beef and pork prices in logs and levels.

| | Pass-through (12 mos.) | | | |
Description	Logs		Levels	
Beef				
Farm price to retail price	0.653	(0.048)	1.058†	(0.115)
Farm price to wholesale price	0.852	(0.031)	0.970†	(0.089)
Farm price to fresh beef retail price	0.547	(0.038)	0.911†	(0.106)
Wholesale price to retail price	0.760	(0.037)	1.013†	(0.100)
Pork				
Farm price to retail price	0.381	(0.058)	0.955†	(0.099)
Farm price to wholesale price	0.550	(0.057)	0.804	(0.063)
Wholesale price to retail price	0.628	(0.071)	0.992†	(0.087)

Note: Long-run pass-through in levels and logs is $\sum_{k=0}^{K} b_k$ from specifications (2) and (3), using a horizon of $K = 12$ months. Newey-West standard errors in parentheses. † indicates that an estimate is statistically indistinguishable from one.

To do so, I subtract monthly corn gluten feed and corn gluten meal prices (also collected by the USDA) from the corn farm price.

Figure E2 shows time series of the corn farm price, the corn farm price after correcting for byproducts, and four downstream products. Notably, the downstream products exhibit considerable differences in how large the cost of corn is as a share of the product price as well as how sticky the output price appears to be. For example, corn prices constitute a large share of the total price of corn starch and corn meal, two products which exhibit high price flexibility, but a smaller share of the price of corn syrup and dextrose, which both exhibit more rigid prices.

Figure E3 plots the pass-through in levels at the six-month and twelve-month horizons from corn farm prices to corn market prices and downstream corn products. At the six month horizon, pass-through remains incomplete for a number of downstream products, especially those with lower corn cost shares. However, by twelve months, pass-through in levels is indistinguishable from one for the majority of the products.

Figure E4 illustrates how pass-through in levels and in logs varies for these downstream products at different horizons. Pass-through in levels at short horizons is positively correlated with the cost share of corn, but flattens and approaches one for all products at longer time horizons. Meanwhile, log pass-through is strongly correlated with the cost share of corn at all horizons.

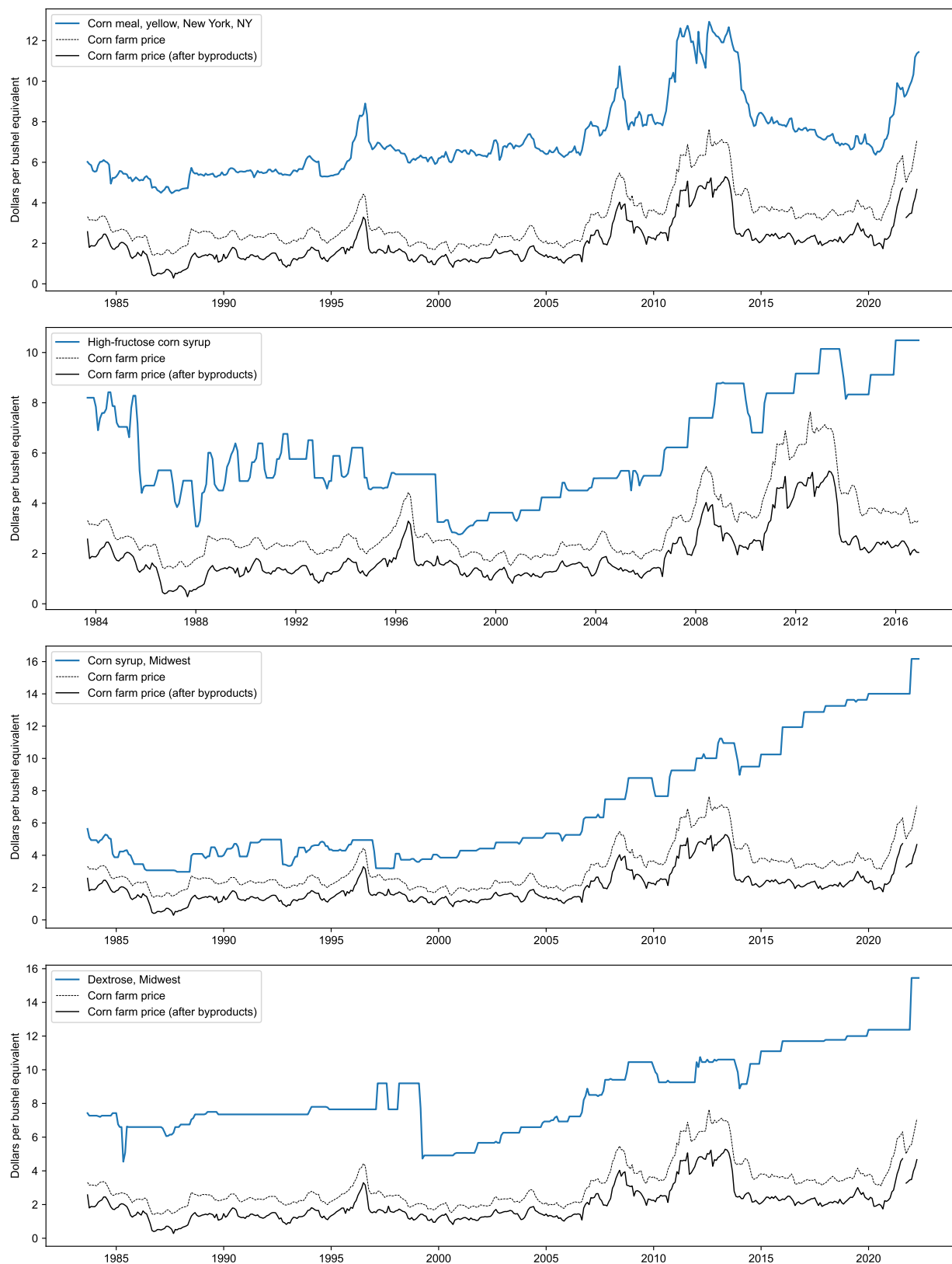**Figure E2:** Price series for corn and four downstream products.

Table E2: Corn and downstream products.

Commodity	Average share of corn (farm price) as % of price
Corn markets:	
No. 2 yellow, Minneapolis, MN	101
No. 2 yellow, Omaha, NE	100
No. 2 yellow, Central IL	99
No. 2 yellow, Toledo, OH	96
No. 2 yellow, Kansas City, MO	95
No. 2 yellow, Chicago, IL	94
No. 2 yellow, St Louis, MO	93
No. 2 yellow, Memphis, TN	91
No. 2 yellow, Gulf ports, LA	85
Corn products:	
Corn starch, Midwest	50
Corn meal, Chicago	33
High-fructose corn syrup	31
Corn syrup, Midwest	31
Corn meal, New York	27
Dextrose, Midwest	23

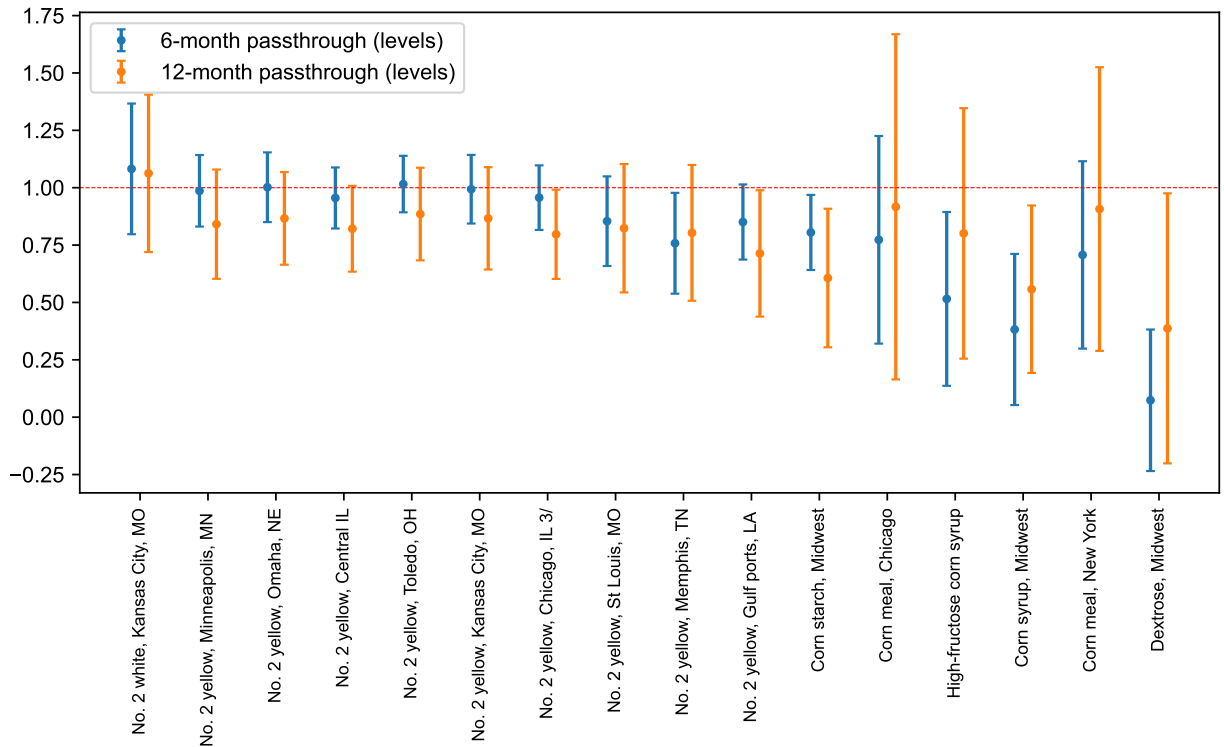**Figure E3:** Pass-through of corn farm price increases in levels.

Figure E4: Pass-through of corn farm price to downstream products, in logs and levels.