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Working paper

24/36

Cheapflation and the rise of inflation inequality

Economic and Social Research Council

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August, 2024

Abstract

The period 2021-2023 saw prices rising at historically high rates. Using household scanner data for fast-moving consumer goods, we show that this was accompanied by historically high rates of inflation inequality. We document systematic increases in the relative prices of cheaper product varieties ('cheapflation') over this period, and show that this drove differences in household-specific inflation rates. After accounting for substitution effects and adjusting for non-homothetic preferences, we show that the inflation-income gradient translated into a similar gradient in the cost of maintaining pre shock living standards. We also show evidence, using an earlier exchange rate shock that followed the UK's vote to leave the European Union, that differential proportional pass-through of cost shocks is likely to be a factor driving cheapflation.

Keywords: inflation, cost-of-living, inequality, heterogeneity, non-homotheticity

JEL classification: D12, D30, E31, I30

Acknowledgements: The authors would like to gratefully acknowledge financial support from the ESRC under grant number CPP: ES/T014334/1. Chen also gratefully acknowledges financial support from the UBEL DTP Research Training Support Grant. We also would like to thank Richard Blundell, Rachel Griffith, Xavier Jaravel, Lars Nesheim, Áureo de Paula and Imran Rasul for helpful comments. The contents of this publication are the authors' own analysis and findings. The analysis and findings have been undertaken by the authors on the basis of their own work using data supplied by the Kantar Take Home Purchase Panel (all use is subject to Kantar UK Limited's terms and conditions) and Kantar does not represent any endorsement of the efficacy or accuracy of authors' analysis and findings. All errors and omissions remain the responsibility of the authors of this publication. Correspondence: tao.chen.16@ucl.ac.uk, peter_l@ifs.org.uk, moconnell9@wisc.edu

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1 Introduction

When consumption baskets differ by income, unequal price rises can lead to inequalities in inflation rates between rich and poor. Economists since at least Marshall (1926) have recognised that these can reinforce inequalities in income growth.[1] For this reason, various statistical agencies now publish separate inflation rates by income or other groups, reflecting differences in spending across categories of goods measured in household budget surveys.[2] However, such measures do not take account of differences in price changes *within* categories of goods, despite evidence that aggregation biases can obscure inflation inequality even in normal times (Kaplan and Schulhofer-Wohl (2017), Jaravel (2019)), and that cost shocks can lead to proportional price changes that differ within categories across similar products (Hong and Li (2017)). These biases might be particularly severe in times of high inflation, which is precisely when estimates of the effect of price rises on different groups are most in demand.

In this paper we study the extent and causes of inflation inequality during the 2021-2023 "cost-of-living crisis", which saw inflation rates rise to levels not seen in rich countries for more than three decades. We use micro data that covers purchases by household in Great Britain of fast-moving consumer goods, one of the sectors that recorded the highest general price rise.[3] The dataset tracks households through time, allowing us to build household-specific inflation and cost-of-living measures, and contains barcode (or UPC) level price and expenditures, meaning we observe price dynamics and consumption baskets free of aggregation. We show that the period 2021-2023 is characterised by a high and, in recent times, unprecedented degree of inflation inequality, almost entirely accounted for by differences in consumption baskets across similar products within narrow categories, and driven by cheapflation – systematically larger price increases for cheaper goods most favoured by the poor. This pattern persists after we take account of substitution responses (people lowering their exposure to price rises by switching to lower priced alternatives) and strip-out the effects of preference non-homotheticities (budget share reallocation

[1]Marshall writes "The same change of prices affects the purchasing power of money to different persons in different ways. For to him who can seldom afford to have meat, a fall of one-fourth in the price of meat accompanied by a rise of one-fourth in that of bread means a fall in the purchasing power of money; his wages will not go so far as before. While to his richer neighbour, who spends twice as much on meat as on bread, the change acts the other way."

[2]This includes the US Bureau of Labor Statistics (Klick and Stockburger (2021, 2024)) and the UK Office for National Statistics (Office for National Statistics (2022)).

[3]In the UK Consumer Price Index (CPI) inflation peaked at an annualised rate of 11.1% in October 2022, its highest rate in over 40 years, and the food and drink CPI index peaked at an annualised rate of increase of 19.1% in March 2023.

due to falling living standards), meaning that inflation inequality drove inequality in the cost of maintaining a fixed standard of living. We also show evidence, using an earlier cost shock associated with the UK's vote to leave the European Union, that a driver of cheapflation (and hence cost-of-living inequality) is larger proportional pass-through of cost shocks for low quality goods.

We begin by documenting that the nine quarter period of elevated price growth over 2021Q3-2023Q3 led to a large and unprecedented degree of inflation inequality. Specifically, over this period average cumulative inflation was 26.6% and the standard deviation across households was 5.6%. To capture inequality in inflation, we bin households into fixed percentiles based on their annual expenditure (to proxy for permanent income), and show that households in the top decile of the expenditure distribution experienced inflation that was 7.7 percentage points lower than those in the bottom decile. This pattern holds accounting for differences in prices paid for the same goods (which act to reinforce inflation inequality), and if we measure income using banded current household income. The inflation income gradient is unprecedented in the UK in recent years; none of the nine quarter periods over 2012-2020 show comparable differences between the inflation rates of rich and poor.[4]

The inflation inequality could be driven by variation in consumption baskets across broad product groups (e.g., alcohol vs. dairy), product categories (e.g., beer vs. wine), or across narrow products within product categories. Using a layered price index, that allows us to decompose the contribution made to inflation inequality due to heterogeneity in spending shares at these three levels, we show that it is differences within product categories that is responsible for 54.5% of inflation dispersion across households and that is entirely responsible for driving the inflation gradient. This points towards a significant aggregation bias in measures of inflation inequality using data from household budget surveys, which typically only allow for a limited degree of product disaggregation.

To better understand why within product category variation in consumption baskets drives inflation inequality, we segment the product space into a 10 rung *quality ladder*, based on within-product category differences in unit prices. We show price growth is much stronger, in proportional terms, for products on lower

[4]Previous research, using data that is aggregated over narrow products, has found quantitatively very small differences in inflation rate for poor and rich households in the UK over 1976-2014 (Crawford and Oldfield (2002), Leicester et al. (2008), Adams and Levell (2014)). The literature using granular data to study inflation inequality in the US documents a gap in average annual inflation rates over 2004-2015 between the bottom and top income quintiles of 0.35 percentage points (Jaravel (2019)) and a gap between the bottom and top income quartiles over 2008-2013 – when US Consumer Price Index inflation peaked at 3.8% in 2008 – of 0.85 percentage points (Argente and Lee (2021)).

rungs of the quality ladder; for instance, over 2021Q3-2023Q3 products on the bottom two rungs exhibit average price rises of 34.2%, whereas those on the top two rungs have average price rises of 18.3%, a pattern known as *cheapflation*.[5] As worse-off households tend to allocate a disproportionate share of their spending to lower quality products, they are more exposed to these price rises, and experience higher inflation rates. Comparing the recent inflation spike to previous periods, we show that it is the pattern of price dynamics across the quality ladder (rather than changes in the make-up of consumption baskets across households of different resources) that is novel and therefore that is key to explaining the recent inflation inequality.

We use a Laspeyres price index to document inflation inequality as this is the index number underlying the official Consumer Price Index (CPI) and it provides a transparent exact decomposition of inflation inequality into the contribution from product group, category and within-category heterogeneity in consumption baskets. However, this index only translates into an exact cost-of-living index (i.e., the change in the cost to a household of maintaining a fixed standard of living) if households do not substitute across products between two comparison periods. Given the pronounced pattern of relative price changes across the quality ladder, it is likely that households did respond by reallocating their expenditure. The standard solution to substitution bias is to use a superlative index, which provides a second-order approximation to a cost-of-living index for arbitrary homothetic preferences (Diewert (1976)). A superlative Törnqvist index suggests household substitution responses act to lower cost-of-living increases by 1.6 percentage points on average, and by 2.6 percentage points for households in the bottom decile of the expenditure distribution. In other words, it suggests that substitution effects acted to mitigate, but not overturn inflation inequality. We also show that using the standard Feenstra (1994) correction for bias arising from the product entry and exit has little impact on the size of the inflation gradient.

Superlative price indexes can provide a biased picture of cost-of-living changes when preferences are non-homothetic. A cost-of-living index measures how the expenditure needed to maintain a fixed standard of living varies over time. If in reality households' living standards are falling over time, and if this leads them to reallocate their spending through an income effect, then superlative indexes may no longer approximate a cost-of-living index well, as they are based on observed rather than constant-utility compensated budget shares (Samuelson and Swamy

[5]The origins of this term are recent. From the second half of 2023 it was used by some media outlets (especially in France and Belgium). We use the term to describe a pattern of higher price growth for initially cheaper products, consistent with Cavallo and Kryvtsov (2024).

(1974), Theil (1976)). We show evidence that part of households' responses to the 2021-2023 cost-of-living crises likely reflected an income effect associated with falling living standards. First, we show that, on average, households reallocated their spending to products at the top and bottom of the quality ladder. The switch towards products at the top, which saw less price growth, is consistent with substitution behaviour. The switch towards products at the bottom of the quality ladder, despite these products becoming relatively more expensive, is consistent with either an income effect (trading down to products that are cheaper in absolute terms as living standard fall) or these goods having inelastic own-price demands. Second, we show that the share of spending that households allocated to products at different points of the quality ladder is strongly related to their total spending. For instance households at the 10^{th} percentile of the expenditure distribution allocated around 15.4% (7.2%) of their total budget to products in the bottom (top) decile of the quality distribution, while those at the 90^{th} percentile allocated 6.8% (12.6%). These patterns point towards lower quality products being necessities and higher quality product luxuries. Lastly, we show that *within* household changes in spending over 2021-2023 are consistent with households moving along Engel curves; with, for instance, those households with the biggest spending falls raising their low quality product spending shares the most and their high quality spending share the least.

A recent literature (Baqaee et al. (2024), Jaravel and Lashkari (2024)) has developed methods for building flexible approximations to a cost-of-living index whilst accommodating non-homothetic preferences.[6] We use the method developed by Jaravel and Lashkari (2024) to correct the Törnqvist index for preference non-homotheticities. The method recursively builds up a sequence of household-specific utility indices, measuring the cost of achieving the household's realised utility level in constant £s by using an estimate of the slope of the cross-sectional relationship between utility and inflation rates to capture non-homotheticites (e.g., a household's falling living standards driving up their inflation rate as they switch toward low quality, high price growth, products due to an income effect). The corresponding approximation to the non-homothetic cost-of-living index can be recovered by the difference in the current period cost of achieving realised utility (i.e., nominal expenditure) and the constant £ cost. We find non-homotheticity bias in the Törnqvist index, when used as a measure of changes in the cost of maintaining living standards at their original 2021Q3 level, ranges from 0 for the poorest households,

[6] This literature also includes a set of papers that account for non-homotheticities by specifying a particular form of non-homothetic preferences (e.g., Fajgelbaum and Khandelwal (2016), Comin et al. (2021), Handbury (2021), Atkin et al. (2024)).

to over 1 percentage point for the richest. The gradient in non-homotheticity bias partially counteracts the gradient in substitution effects suggested by the unadjusted Törnqvist index. In other words, for better-off households the Törnqvist index understates the extent of substitution effects, as it is contaminated by these households switching to products with high price growth through an income effect. Overall, after accounting for substitution effects and non-homotheticities, we find that households in the top decile of the expenditure distribution experienced cost-of-living rises that were 7.2 to percentage points lower than those in the bottom decile.

Inflation and cost-of-living inequality was driven by cheaper products exhibiting relatively high price growth. As a central driver of these price increases was cost increases (see Competition and Markets Authority (2023), who document input prices increases, including for ingredients, fertiliser, labour, energy, and packaging), differential pass-through of cost shocks is likely to have played a key role in driving cheapflation. While the benchmark model of CES demands and monopolistic pricing (Dixit and Stiglitz (1977)) implies complete proportional cost pass-through, a number of papers (e.g., Gopinath and Itskhoki (2010), De Loecker et al. (2016), Mrázová and Neary (2017), Miravete et al. (2018)) have documented empirically that cost pass-through is incomplete in proportional terms and varies across products, and is often approximately one-for-one in level terms (e.g., Nakamura and Zerom (2010), Butters et al. (2022), Sangani (2023)). The 2021-2023 inflationary episode is the result of a complex set of shocks to commodity and energy prices, labour markets, and global supply chains. Therefore, to assess evidence of differential proportional pass-through of marginal cost shocks, we use an earlier cost shock driven by the sterling exchange rate depreciation that occurred around the UK's vote to leave the European Union. This provides a relatively clean cost shock that is likely to have had a similar proportional impact on the marginal costs of supplying Eurozone imports, with no direct effect on the costs of domestic products. We show that in the year following the 20 percent depreciation in the £-€ exchange rate, the price of EU products in the lower half of the pre-shock price distribution rose by around 7 percentage points relative to both domestic and higher priced EU produce. This provides the first evidence of Brexit driving relative price rises for cheaper products and points to a pattern of higher proportional pass-through to lower quality products likely being an important contributor to cheapflation and inflation inequality.

Our work belongs to a literature on the measurement of inflation inequality. While early papers found little evidence of inflation rates differing between households belonging to different demographic groups (e.g., Michael (1979) and Hagemann (1982)), more recent work that uses granular micro data has documented inflation inequality during times of moderate aggregate inflation (e.g., Broda and Weinstein (2010), Kaplan and Schulhofer-Wohl (2017), and Argente and Lee (2021)). Jaravel (2019) shows that this inflation inequality can be explained by better-off households benefiting more from product entry and heightened competitive pressures driven by innovation. Our primary contribution is to show that inflation inequality is significantly heightened during the recent period of high aggregate inflation and that it has a novel cause; cheapflation. Cavallo and Kryvtsov (2024) use data on posted prices to show evidence of cheapflation across 10 countries, while Sangani (2023) uses the historic relationship between commodity price shocks and inflation in the US to predict recent increased inflation inequality. This points towards the phenomenon of cheapflation driven inflation inequality, that we document in the UK, being widespread internationally.

We also contribute to a recent literature on adjusting index numbers to account for preference non-homotheticities. This literature has shown that non-adjusted standard- and cost-of-living indexes computed over several decades can suffer from substantial bias (Baqaee et al. (2024), Jaravel and Lashkari (2024)). We show that a combination of falling purchasing power and cheapflation led to expenditure switching across the quality ladder due to both substitution responses to relative prices changes, and households moving along product-level Engel curves, and therefore that non-homotheticity bias can be substantial even over a short time horizon, when prices are rising rapidly.

Finally, our work also contributes to a growing literature studying the welfare effects of the inflationary period following the COVID-19 pandemic (e.g., Pallotti et al. (2023), Afrouzi et al. (2024)) and a literature measuring the impact of the UK's departure from the European Union on consumers (Bakker et al. (2022), Breinlich et al. (2022)).

The remainder of this paper is structured as follows. Section 2 discusses the household scanner data we use. Section 3 documents inflation inequality over the period of high inflation and compares it to previous periods. Section 4 decomposes household-level inflation into the contribution made by across and within product category variation in consumption baskets, and shows the role played by cheapflation in driving inflation inequality. Section 5 documents the extent of inequality in cost-of-living rises after accounting for substitution effects, product entry and preference non-homotheticities. Section 6 presents our results on the unequal pass-through of the sterling depreciation surrounding the Brexit referendum. A final section concludes.

2 Data

Scanner data. We use household level scanner data that is collected by the market research firm Kantar's Take Home Purchase Panel. The data cover purchases of fast-moving consumer goods (food, alcohol and non-alcoholic drinks, toiletries, cleaning products, and pet foods) brought into the home by a sample of households living in Great Britain (i.e., the UK excluding Northern Ireland). The data cover both purchases made in brick and mortar stores and those made online.

Participating households are in the dataset for many months on average. Each household records all UPCs (or barcodes) that they purchase using a handheld scanner or mobile phone app, and they send their receipts (either electronically or by post) to Kantar. For each transaction we observe quantity, expenditure, price paid and UPCs characteristics. We also observe socio-demographic characteristics of households, including household structure and (banded) income.

Our main period of interest are the nine calendar quarters from 2021Q3 to 2023Q3, which is the period of elevated inflation. For instance, over this period the CPI index for food and non-alcoholic beverages (which, in addition to alcohol and household goods, comprises fast-moving consumer goods), following many quarters of stability, grew from 103.3 to 133.5.[7] Our analysis sample comprises 19,030 households that record their purchases in each year-quarter over this period.[8] We compare inflation inequality over 2021Q3-2023Q3 to four earlier 9 quarter periods, 2012Q1-2014Q1, 2014Q1-2016Q1, 2016Q1-2018Q1 and 2018Q1-2020Q1.[9]

Our data cover c. 200,000 unique UPCs over 2021Q3-2023Q3. As sometimes one UPC is replaced with a new one that is almost indistinguishable, we define products based on the slightly more aggregated combination of brand and size. Kantar

[7] In January 2019 the index was 102.6 and in April 2024 is was 135.6. One exception during the preceding period of price stability was a spike and subsequent reduction in inflation at the beginning of the COVID-19 pandemic associated with a reduction and subsequent recovery in the level of promotion activity (Jaravel and O'Connell (2020a, 2020b)).

[8] We omit from our analysis any households that are not continuously present throughout the 9 quarter period or that ever report quarterly expenditure below the 5^{th} percentile (£114 on average) of the quarterly expenditure distribution. Our results are not sensitive to these restrictions.

[9] We observe 19,000-20,000 household in each of these 9-quarter periods. We omit 2020Q2-2021Q2 from our analysis as purchase patterns are likely to be atypical due to the influence of lockdown and social distancing measures aimed at curbing the spread of COVID-19. However, inclusion of this period in our set of comparison periods makes no material difference to any of our results.

provides extremely disaggregate brand information, and therefore this entails essentially no aggregation over meaningfully different products. Over 2021Q3-2023Q3 there are approximately 90,000 brand-size pairs. Throughout we refer to these as products.[10] We make use of a hierarchical classification of products that is based on information provided by Kantar and that is designed to allocate products into well-defined consumer markets. In particular, products are grouped into 10 segments (bakery, dairy, fresh fruit and vegetables, meat and fish, prepared food, cupboard ingredients, confectionery, non-alcoholic drinks, alcohol and household goods) and 238 categories. For instance, the product *Coca Cola 2 liter bottle* belongs to the category *colas* and the segment *non-alcoholic drinks*.

For fresh fruit and vegetables, fresh meat, cheese and honey Kantar provide information on the products' country of origin, which enables us to categorise whether products are domestic, are imported from the EU or are imported from non-EU countries. We use data covering the period 2015-2017 to quantify the impact of depreciation of the pound around the Brexit referendum on price dynamics. In total the categories with country of origin information cover 17.8% of total fast-moving consumer good spending over this time period.[11]

Our dataset offers several advantages for measuring inflation inequality, relative to other commonly used alternatives. First, it contains up-to-date expenditures, allowing us to construct inflation measures based on up-to-date spending patterns. Second, it tracks households through time, meaning we can construct household-specific inflation rates. Third, expenditure information is recorded at the UPC level, meaning our inflation measures reflect spending patterns across narrowly defined product categories.[12] The data is more detailed than the budget surveys and price microdata that are often used to study household-level inflation inequality. The UK Office for National Statistics publishes estimates of inflation rates across quintiles of the household income distribution, which uses spending data from the Living Costs and Food Survey along with prices for CPI 'classes', of which there are around 100, across all categories of spending (down to the level of food items such as 'bread and cereals', 'meat' etc.) (Office for National Statistics (2022)). Similarly, estimates of household-level inflation from the US using the Consumer Expenditure Survey use

[10] For some fresh produce, such as fruit and vegetables and meat, UPC do not have a well defined brand. In this case we define products by their UPC. All of our results hold if we define all products based on UPCs.

[11] Within categories for which the variable country of origin is available, the reported country of origin is "Not stated" for products accounting for 30% of total spending. We omit these from our analysis.

[12] Store-level scanner data, which now is being used in some countries for CPI construction, also contain disaggregate and up-to-date expenditure information, but as it is store rather than household level it is not appropriate for studying the distribution of inflation across households.

data on budget shares across a few hundred Universal Commodity Codes can are linked to around 300 CPI subindices (Mejia and Hartley (2022)). While these data sources have the advantage, relative to our data, or covering a broader proportion of households' total spending, they are not suitable for measuring the role played by differential price increases across similar products.

Income measure. Our data offer two alternative ways to measure income, either using a measure of current household income (which is reported in £10,000 bands, top-coded at £70,000), or using total household fast-moving consumer good expenditure. Throughout the paper we use total household expenditure, since it is not banded and arguably better reflects permanent income than current income does. However, in the Appendix, we show our main results remain if we instead use the banded household income measure. For both total expenditure and banded income (based on the band midpoint) we equivalize using the OECD modified scale (see Hagenaars et al. (1994)).[13] In each nine quarter period we use equalized expenditure over the initial calendar year of the period to group households into 100 and 10 fixed bins based on equalized expenditure percentiles or deciles. Throughout we refer to these as expenditure percentiles (or deciles).

The quality ladder. It is useful to segment the product space using a more disaggregate classification than product category. We do this by splitting products (within the narrowly defined categories) into "rungs" of the quality ladder using a procedure similar in spirit to previous work (Jaravel (2019); Argente and Lee (2021)) based on differences in products' unit price.

For each product category, over 2021Q3-2023Q3, we estimate the expenditure-weighted regression:

$$P_{it} = \chi_i + \tau_{c(i)t} + \epsilon_{it}, \tag{2.1}$$

where P_{it} is the average quarterly price (per unit volume) paid for product i in quarter t,[14] $\tau_{c(j)t}$ are year-quarter effects, and χ_i are product fixed effects.

To assign products to quality rungs, we first use the set of products available in the four quarters 2021Q3-2022Q2, construct product-level average price over these four quarters, and compute category-specific decile-boundaries for the expenditure weighted-distribution of average prices. For all products we then predict their 2021Q3-2022Q2 average price using $\tilde{P}_{it} = \hat{\chi}_i + \frac{1}{4}\sum_{t=1}^{4}\hat{\tau}_{c(i)t}$, and assign them to

[13] This entails constructing a household equivalized size, with the first adult counting 1, any additional individuals aged 14 or over counting 0.5 and each child under 14 counting 0.3

[14] We compute this as the sum of total expenditure on the product divided by total quantity.

10 rungs of the quality ladder based on how their predicted price compares with the decile-boundaries. This procedure ensures that we assign products that were introduced after 2022Q2 to a quality rung based on a price that adjusts for category-specific price growth. We construct the quality ladder analogously for the four earlier comparison periods.

A potential concern with interpreting differences in unit prices for similar good as reflecting their quality, is that nonlinear pricing can mean the unit price of a small pack of a brand can be significantly higher than a large pack of the same brand (for instance, see Griffith et al. (2009)). We therefore also define the quality ladder based on a procedure that adjusts for nonlinear pricing (see Appendix B) and we show that our results hold under this alternative.

3 Inflation inequality

Measurement. As our baseline measure of household inflation, we use a household-level Laspeyres index. Specifically, let x_{hit} denote household h's year-quarter t spending on product i, which has price P_{it}, and let $s_{hit} = \frac{x_{hit}}{\sum_j x_{hjt}}$ denote the household-specific product i spending share. The (direct) household-level inflation index between two periods s and $t > s$ is:

$$\Pi_{h,(s,t)} \equiv \sum_i s_{his} \left(\frac{P_{it}}{P_{is}} \right) \tag{3.1}$$

Inflation between two non-consecutive periods, for instance period 1 and T, can either be computed by the direct (D) or chained (C) comparison:

$$\Pi^D_{h,(1,T)} = \Pi_{h,(1,T)}$$
$$\Pi^C_{h,(1,T)} = \prod_{t=1}^{T-1} \Pi_{h,(t,t+1)}.$$

Chained comparisons have the advantage that they use up-to-date expenditure weights, allowing for *between* time period spending reallocation. In contrast, the direct comparison uses expenditure weights that are increasingly out dated over time. However, a risk with high frequency chained indices is that they can suffer from chain-drift bias, which can arise if expenditure weights in one period are correlated with price changes in another period (see Ivancic et al. (2011), Diewert (2022)).[15] Unless stated otherwise, throughout the paper we report cumulative inflation based

[15] An index subject to chain drift will fail the multiperiod identity test – if all prices and weights return to their initial period values the chained price index will not in general equal its initial value.

on an index that is chained annually across calendar quarters.[16] We show our results are robust to using a quarter-to-quarter chaining and direct comparisons.

The Laspeyres index is commonly used, forms the basis for official CPI inflation measurement, and can be additively decomposed into the influence of variation in expenditure shares across segments, categories and individual products. However, when interpreted as a measure of a Konüs (1939) exact cost-of-living index, it may suffer from a series of biases (which we address in Section 5).

Our baseline inflation measure uses prices that are common across households. One reason for this is that comparisons in the price index (3.1) can involve the price of a product that a household does not purchase in one of the comparison periods. In the context of the UK grocery industry, where retailers have national store coverage and practice national pricing policies (see Competition Commission (2008)), the assumption of common prices is arguably more innocuous than in other settings. Nonetheless, there is likely to be some price dispersion for identical products, due for instance, to discounts that are in place for only a limited number of weeks and only in some retailers. If households of different incomes alter their propensity to take advantage of low prices over time this could impact inflation inequality. To assess this we use a price index suggested in Aguiar and Hurst (2007), which measures dispersion in price paid for a *fixed* basket of products.

Let q_{hit} denote the volume of product i purchased by household h in year-quarter t and define the household-specific price by $p_{hit} = x_{hit}/q_{hit}$.[17] Had the household paid average prices for their basket their expenditure would have been $\tilde{x}_{ht} = \sum_i q_{hit} P_{it}$. The Aguiar and Hurst (AH) index compares the true cost of the household's basket ($x_{ht} = \sum_i x_{hit}$) with its cost at average prices:

$$\Pi^{AH}_{ht} = \frac{x_{ht}}{\tilde{x}_{ht}}. \tag{3.2}$$

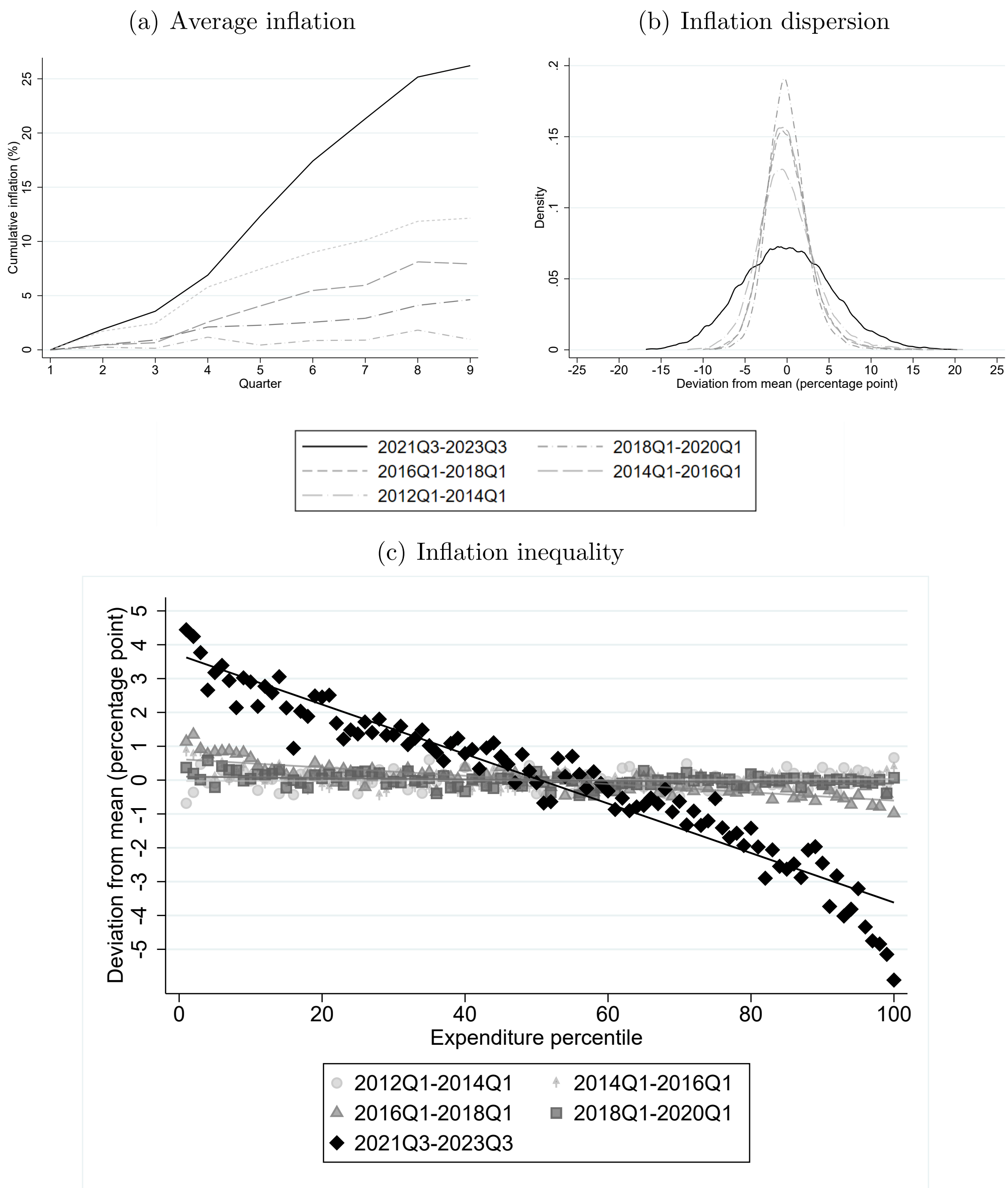
Results. In Figure 1(a) we show average household-level cumulative inflation across the nine quarters 2021Q3-2023Q3, and across the earlier nine quarter periods. It shows that inflation was historically high over the recent period, with a cumulative increase in excess of 25%. This compares to an increase of 13% over the

A common reason why chain drift arises is that households engage in stockpiling, so that a product purchased in one period may be consumed in latter periods.

[16]For instance, cumulative inflation over 2021Q3-2023Q3 is based on $\Pi_{h,(1,2)} \times \Pi_{h,(2,3)}$, where $t = 1$ is 2021Q3, $t = 2$ is 2022Q3, and $t = 3$ is 2022Q3. This allows for updating in expenditure shares, while sidestepping chain drift that can arises from seasonal patterns in prices and expenditures.

[17]Note that the common price and household-specific prices are related by: $P_{it} = \frac{\sum_h x_{hit}}{\sum_h q_{hit}} = \sum_h \frac{q_{hit}}{\sum_{h'} q_{h'it}} p_{hit}$.

Figure 1: *Household-level inflation*



Notes: Authors' calculations using Kantar's Take Home panel (2012-2023). Panel (a) shows a plutocratic average (across households) of cumulative inflation for each of the 9 quarters using a quarter-to-quarter chained Laspeyres index. Panel (b) show a kernel density estimate of the distribution of 9^{th} quarter cumulative inflation. Panel (c) plots the relationship between 9^{th} quarter cumulative inflation and percentile of the expenditure distribution a household belongs to; there is a marker for each percentile and a line of best fit. We allocate households to expenditure percentiles based on their equivalized spending over the initial calendar year of the relevant nine quarter period. Cumulative inflation in panels (b) and (c) is computed using an annually chained Laspeyres index.

period 2012Q1-2014Q1, which was the nine quarter period with the second-highest rate of inflation. Panel (b) shows the distribution of household-level cumulative inflation rates across the various nine quarter periods. The elevated average rate of inflation in 2021Q3-2023Q3 is associated with higher dispersion in inflation rates across households. Panel (c) shows the relationship between household-level cumulative inflation and expenditure percentiles. It makes clear that over 2021Q3-2023Q3 there were historically large differences in inflation rates across households of different expenditure levels. While for much of the previous decade or so inflation rates have exhibited at most a modest gradient with expenditure, over the period 2021Q3-2023Q3, the gradient is much larger; for instance, households in the 1^{st} quartile of the expenditure distribution experienced inflation rates that were 5.6 percentage points higher than those in the 4^{th} quartile.

In Figure 2 we summarise the evolution of the AH index over time and show the implications for inflation inequality. In panels (a) and (b) we summarise how the AH index varies over 2012-2023, across quartiles of the equivalized expenditure distribution. To do this, in each calendar year (2012-2023), we group households into quartiles of the expenditure distribution and we report the average AH index across households in each expenditure quartile (expressed as a deviation from the quarter mean across all households) in each year-quarter. Panel (a) shows the index when products are defined based on brand-size. Panel (b) show the index when we re-define products based on the combination of brand-size and retailer (which is relevant for branded products sold by multiple retailers, but leaves the definition of private label products unaffected). The lines in panel (a) reflect the influence of price dispersion across and within retailers; those in panel (b) strip out the former and therefore reflect only within retailer price dispersion.

In 2012 households in the top expenditure quartile paid around 2 percentage points more for a fixed basket of goods than those in the bottom quartile. Around 1 percentage point was due to cross retailer variation and 1 percentage point due to within retailer variation. This gap has closed over time; by 2023 those in the top expenditure quartile paid only around 1.1 percentage point more than those in the bottom quartile (split approximately evenly between the influence of across and within retailer dispersion). This decline, implies differences in price paid for identical goods act to reinforce inflation inequality.

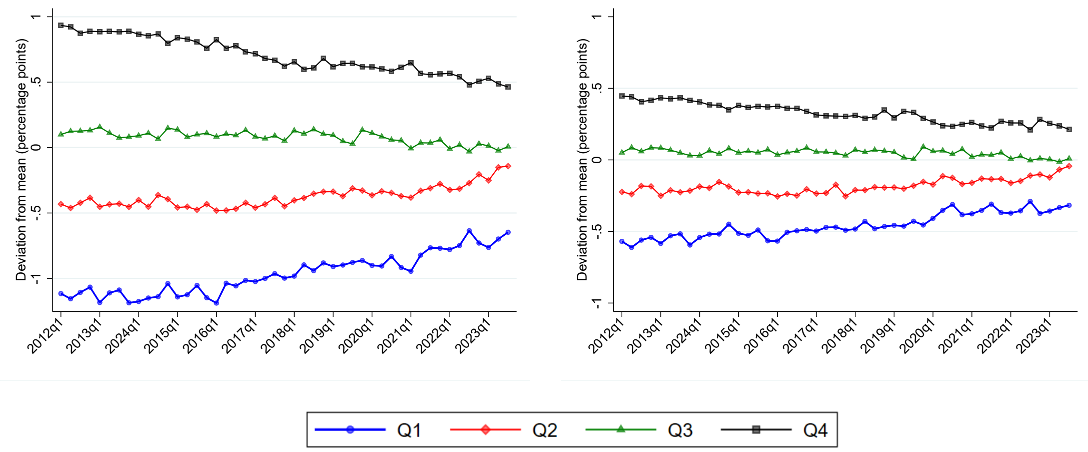
In panel (c) of Figure 2 we confirm that this is the case by showing the relationship between household-level cumulative inflation over 2021Q3-2023Q3 based on common prices (the blue line and markers; repeating information in Figure 1(c)) and based on price computed at the expenditure quartile level [18] (the red line and markers). It shows that the inflation gradient is (to a modest degree) larger in the

[18] Specifically, we compute the expenditure quartile-specific price for (i,t) as $P_{it}^{r} = \frac{\sum_{h \in \mathcal{R}_r} x_{hit}}{\sum_{h \in \mathcal{R}_r} q_{hit}}$ where $\mathcal{R}_r$ denotes the set of households that belong to quartile r of the calendar year-specific annual equivalized expenditure distribution.

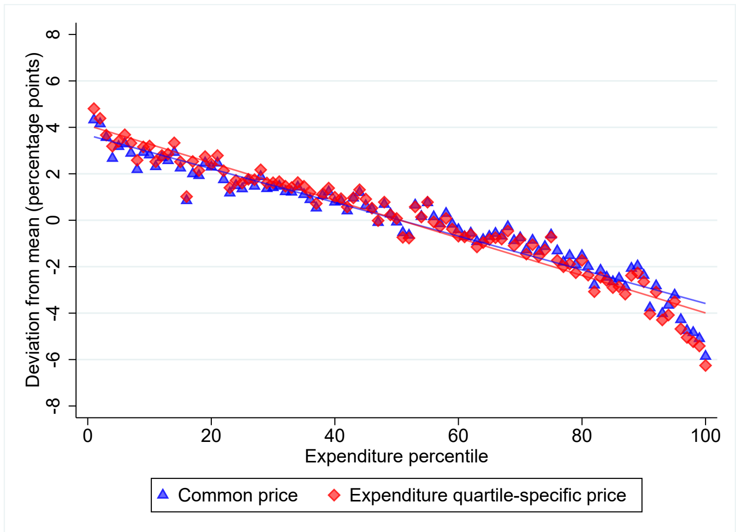
latter case. In sum, variation in price paid appears to slightly reinforce, rather than unwind, our finding of significant recent inflation inequality.

Figure 2: *Price dispersion*

(a) AH index (product) (b) AH index (product-retailer)



(c) Household inflation (2021Q3 - 2023Q3)



Notes: Authors' calculations using Kantar's Take Home panel (2012-2023). Panels (a) and (b) show within expenditure quartile average (across households) AH index at quarterly frequency over 2012Q1-2023Q3. In panel (a) products are based on brand-size and in panel (b) they are based on brand-size and retailer. Panel (c) plots the relationship between cumulative inflation and percentile of the expenditure distribution a household belongs. We allocate households to expenditure percentiles based on their equivalized spending in 2021. Cumulative inflation in panel (c) is computed using an annually chained Laspeyres index.

Robustness. In Appendix C we show that when we instead use banded household income as our measure of how well-off households are, we continue to find that the most recent nine quarter period exhibits a pronounced gradient in inflation inequality that is absent in earlier periods. We also continue to find a closing over time in the gap in prices paid for a fixed basket of goods between those at the top and bottom of the income distribution. In addition, we show that using either a quarter-to-quarter chained or a direct comparison, in place of a calendar quarter to calendar quarter chained index, results in a quantitatively very similar gradient.

4 Sources of inflation inequality

Inflation inequality is driven by a systematic relationship between product-specific price rises and the shares of their consumption baskets that households of different resource levels allocate to products. In this section we show that it is principally variation in consumption baskets across similar products within narrowly defined categories that drives inflation inequality (and therefore data that entails product aggregation will understate the inflation gradient) and that within category differences are due to an inverse relationship between product-level price rises and their position on the quality ladder.

4.1 Aggregation bias

Is variation in household spending shares across broad product segments, product categories, or products within categories primarily responsible for accounting for inflation inequality? To answer this we use a layered (or hierarchical) version of the household-specific price index.

Let Ω^c denote the set of products that belong to product category c and Γ^g denote the set of product categories that belong to segment g, and define the household-specific within product category spending share of product i, s^c_{hit}, within segment spending share of product category c, s^g_{hct}, and spending share of segment g, s_{hgt}, as:

$$s^c_{hit} = \frac{x_{hit}}{\sum_{i' \in \Omega^c} x_{hi't}} \qquad s^g_{hct} = \frac{\sum_{i \in \Omega^c} x_{hit}}{\sum_{c' \in \Gamma^g} \sum_{i \in \Omega^{c'}} x_{hit}} \qquad s_{hgt} = \frac{\sum_{c \in \Gamma^g} \sum_{i \in \Omega^c} x_{hit}}{\sum_{g'} \sum_{c \in \Gamma^{g'}} \sum_{i \in \Omega^c} x_{hit}}.$$

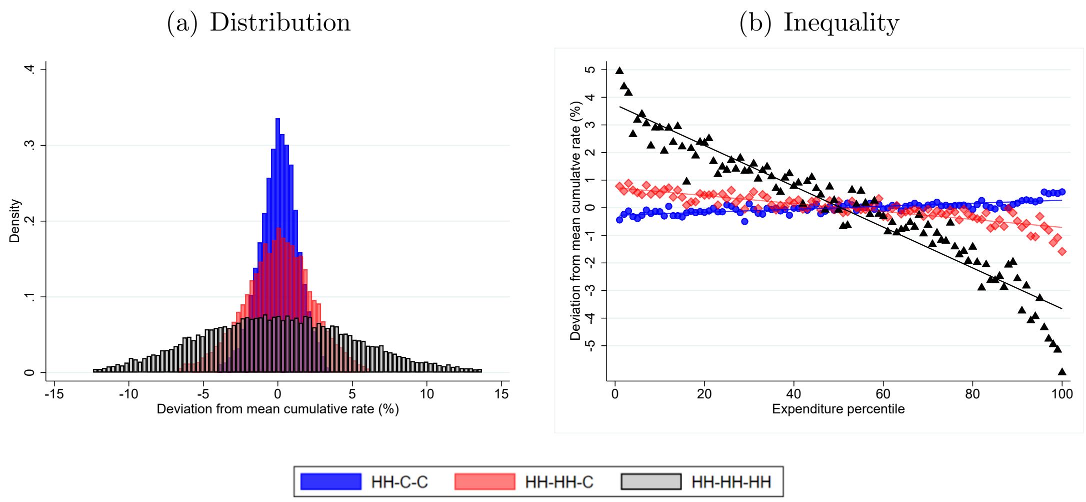
Note that $s_{hit} = s^c_{hit} s^g_{hct} s_{hgt}$. The direct comparison between two periods s and $t > s$ under the hierarchical index takes the form:

$$\begin{aligned} \mathbb{P}_{hc,(s,t)} &= \sum_{j \in \Omega^c} s^c_{his} \left(\frac{P_{it}}{P_{is}} \right) \\ \mathbb{P}_{hg,(s,t)} &= \sum_{c \in \Gamma^g} s^g_{hcs} \left(\mathbb{P}_{hc,(s,t)} \right) \\ \Pi^{Hierarchical}_{h,(s,t)} &= \sum_{g} s_{hgs} \left(\mathbb{P}_{hg,(s,t)} \right). \end{aligned} \tag{4.1}$$

Note that $\Pi^{Hierarchical}_{h,(s,t)} = \Pi_{h,(s,t)}$.[19]

We use the hierarchical index to isolate the roles of differences in spending shares across households at the segment, category and product level in accounting for differences in household-specific inflation. We do this by switching household-specific spending shares with population average spending shares at different levels of the index. For instance, to isolate the importance of heterogeneity in segment shares, we replace the household-specific category and product shares with their across household averages.

Figure 3: *Household-level hierarchical inflation*



Notes: Authors' calculations using Kantar's Take Home panel (2021-2023). Panel (a) show a histogram of the distribution of cumulative inflation over 2021Q3-2023Q3 and panel (b) plots the relationship between cumulative inflation and percentile of the expenditure distribution a household belongs to; there is a marker for each percentile and a line of best fit. HH-C-C uses household specific segment shares and average within-segment shares, HH-HH-C uses household-specific segment and category shares and average within-category shares. HH-HH-HH uses household-specific shares at all levels. We allocate households to expenditure percentiles based on their equivalized spending in 2021. Cumulative inflation is computed using an annually chained Laspeyres index.

In Figure 3 we plot the distribution of household-level cumulative inflation over 2021Q3-2023Q3 (panel (a)) and the inflation gradient (panel (b)), with black lines and markers, replicating information from Figure 1. We also show patterns when we 'switch off' heterogeneity in spending across products within categories (red lines and markers) and additionally switch off across household variation in spending across categories within segments (blue lines and markers). Panel (a) shows that across household variation in spending across products *within* product categories contributes a substantial fraction of the dispersion in household inflation rates, raising the standard deviation from 1.4 (based only on heterogeneity in segment shares) and 2.5 (plus categories) to 5.6 (plus product). Panel (b) shows that heterogeneity

[19]This is a useful property of the Laspeyres index that is referred to in the index number literature as 'consistency in aggregation'.

in spending across products within product categories is *solely* responsible for the gradient in inflation. This emphasises the importance of using detailed household spending and price information to study inflation inequality.

4.2 Cheapflation

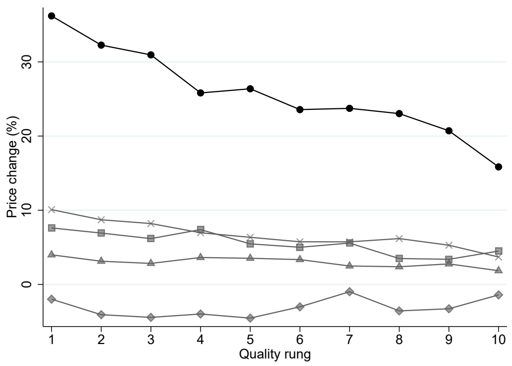
Figure 3(b) makes clear that the inflation gradient over 2021Q3-2023Q3 was driven by variation in price rises and spending shares across products within narrowly defined categories. In Figure 4 we show that this was a consequence of worse off households allocating a relatively high fraction of their consumption baskets to low quality products, and these products exhibiting the strongest price growth. In other words cheapflation drove inflation inequality.

In panel (a) of Figure 4 we report the increase in average price (weighted by initial period aggregate spending shares) of products on each rung of the quality ladder (black line). It shows that price rises (in proportional terms) were substantially higher for products at the bottom of the quality ladder, compared to those towards the top. For instance, for products on the bottom two rungs (1 and 2) the average price increase over 2021Q3-2023Q3 is 36.2% and 32.3%, while those on the top two rungs (9 and 10) it is 20.7% and 15.8%. Panel (a) also shows price growth across the quality ladder in the set of prior nine quarter periods; showing that in these periods of moderate inflation (or mild deflation in 2014Q1-2016Q1) price rises were similar across the quality ladder.

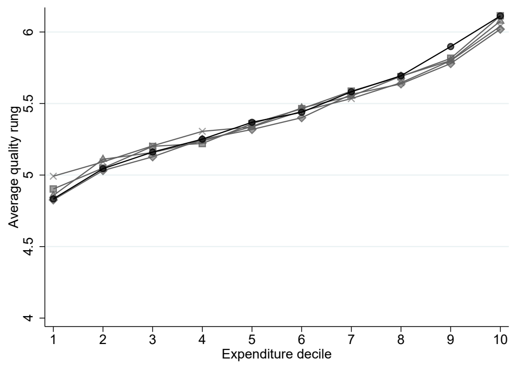
In panel (b) we show how the average quality rung of products that households purchased in the initial quarter of each nine quarter period varies across deciles of the expenditure distribution. It shows that the better off (i.e., households with higher equivalized spending) systematically purchase products that, on average, are of higher quality. This relationship is remarkably stable across all nine quarter periods. Taken together, panels (a) and (b) suggest that it was novel features of price changes, rather than changes in spending patterns, that account for the greater degree of inflation inequality over the recent inflation episode.

Figure 4: *Inflation and quality rungs*

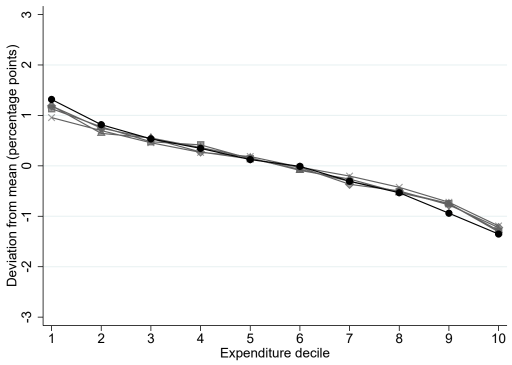
(a) Prices changes across quality rungs



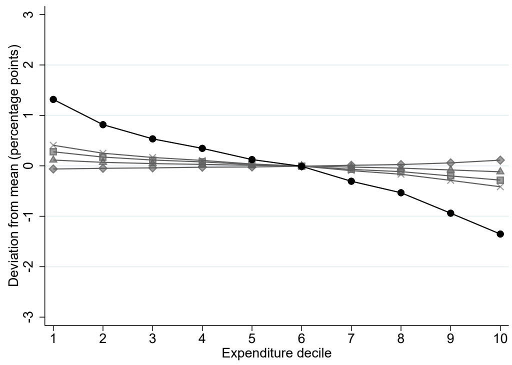
(b) Average quality rung, by expenditure

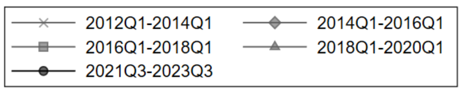


(c) Inflation with recent period price changes



(d) Inflation with recent period budget shares





Notes: Authors' calculations using Kantar's Take Home panel (2012-2023). Panel (a) reports initial quarter-weighted increases in average price over the nine quarter period for products on each rung of the quality ladder. Panel (b) reports the average quality rungs of households' purchases by deciles of the expenditure distribution. Panels (c) and (d) report the relationship between cumulative inflation and expenditure decile in each nine quarter period, calculated using 2021Q3-2023Q3 price changes (panel (c)) and 2021Q3-2023Q3 spending shares (panel (d)). We allocate households to expenditure deciles based on their equivalized spending over the initial calendar year of the relevant nine quarter period. Cumulative inflation is computed using an annually chained Laspeyres index.

To assess whether this is indeed the case we calculate household-specific inflation rates for each period using 2021Q3-2023Q3 price changes (panel (c)) and 2021Q3-2023Q3 spending weights (panel (d)). As there is some product entry and exit over 2012-2023 at the product level, we undertake this exercise by re-defining products based on the interaction of product category and quality ladder rungs. The black lines in panel (c) and (d) use both 2021Q3-2023Q3 prices and shares. They show a similar pattern of inflation inequality as the index computed over brand-size (Figure 1(c)). Therefore inflation computed over aggregated products is able to capture the inflation gradient, as long as the aggregation embeds the quality ladder. The inflation gradient using 2021Q3-2023Q3 price changes but earlier spending shares looks very similar to the realised inflation gradient over this period. Conversely, the gradient based on 2021Q3-2023Q3 shares but earlier price changes is flat. This shows that the recent rise in inflation inequality is driven by differences in the nature of price changes – and in particular, cheapflation – that occurred relative to previous periods.

In Appendix C we replicate Figure 4 when we first adjust prices for within brand non-linear pricing before constructing the quality ladder. The quantitative patterns are very similar.

5 Consumption responses and implications for the cost-of-living

To this point we have computed household-specific inflation rates based on a Laspeyres index. This price index forms the basis of CPI measurement and therefore is a natural benchmark. However, there are a number of reasons why it may give a biased picture of changes in a household's cost-of-living, defined as how the cost of reaching a fixed utility level changes over time. We discuss three biases, arising from substitution effects, new product varieties, and income effects, and show taking account of each does not unwind the pattern of inflation inequality.

The cost-of-living index. Suppose a household has preferences encoded by the indirect utility function, $v_h(\mathbf{P}, x)$, and its inverse, the expenditure function $e_h(\mathbf{P}, u) = v_h^{-1}(\mathbf{P}, .)$, where $\mathbf{P}$ is price vector, x denotes total expenditure and u the corresponding utility level. Consider two periods $\tau = \{s, t\}$ and let $(\mathbf{P}_\tau, x_{h\tau}, u_{h\tau})$ denote the price vector and realised expenditure and utility levels. The cost-of-living index between s and t is defined by $\mathbb{P}_h(\mathbf{P}_s, \mathbf{P}_t; u) = \frac{e_h(\mathbf{P}_t, u)}{e_h(\mathbf{P}_s, u)}$ and measures the change in the minimum expenditure necessary to achieve utility level u. In general the value of the cost-of-living index depends on the reference utility level. An exception is when the household's preferences are homothetic, in which case the composition of their consumption basket (i.e., spending shares) is independent of their level of total expenditure, the expenditure function takes the form $e_h(\mathbf{P}, u) = e_h^H(\mathbf{P})u$ and the cost-of-living index is $\mathbb{P}_h(\mathbf{P}_s, \mathbf{P}_t; u) = \mathbb{P}_h^H(\mathbf{P}_s, \mathbf{P}_t) = \frac{e_h^H(\mathbf{P}_t)}{e_h^H(\mathbf{P}_s)}$.

Substitution bias. The Laspeyres index provides an exact measure of the cost-of-living if preferences are Leontief, meaning the household consumes products in fixed proportions, not altering their choices in response to relative price changes.[20] If households respond by switching away from products with relatively high price increases and towards those with relatively low price increases, the Laspeyres index will suffer from a positive substitution bias, meaning it will overstate cost-of-living increases.

Superlative price indexes provide an approximation to the cost-of-living index, whilst allowing for the possibility of substitution effects. In particular, they provide second-order approximations to an arbitrary *homothetic* cost-of-living index (Diewert (1976)).[21] The assumption of homothetic preferences means any reallocation of spending shares across products is due only to substitution (or relative price) responses, ruling out any spending reallocation due to declining standards of living. One superlative index is the Törnqvist index, which, in log terms takes the form:[22]

$$\log \mathbb{P}^T_{h,(s,t)} = \frac{1}{2} \sum_i \left(s_{his} + s_{hit}\right) \left(\log P_{it} - \log P_{is}\right)$$

In Figure 5(a) we show that the inflation gradient computed using the Törnqvist index and, for purposes of comparison, the Laspeyres index. Panel (b) shows the difference between the two, providing a measure of substitution bias in the Laspeyres index (under the assumption of homothetic preferences). The figure shows that this measure of substitution bias is larger for households towards the bottom of the expenditure distribution. In other words, the superlative index points towards worse-off households lowering their cost-of-living rise to a greater extent than better-off households through reallocating the spending to goods exhibiting lower price

[20] If the household's direct utility function has the Leontief form, $U = \min\{\frac{q_{h1}}{a_{h1}}, \dots, \frac{q_{hI}}{a_{hI}}\}$ where $a_{hi} > 0$ for all i, their expenditure function takes the form $e^H_h(\mathbf{P}) = \sum_i a_{hi} P_i$ and their cost-of-living index takes the from $\mathbb{P}^H_h(\mathbf{P}_s, \mathbf{P}_t) = \frac{\sum_i a_{hi} P_{it}}{\sum_i a_{hi} P_{is}}$. Since (Hicksian) demands satisfy $q_{hi} = a_{hi} u$, $\mathbb{P}^H_h(\mathbf{P}_s, \mathbf{P}_t)$ collapses to the Laspeyres index. Note, although the Laspeyres index rules out substitution effects within comparison periods, chaining the index does allow for consumption basket changes across comparison periods.

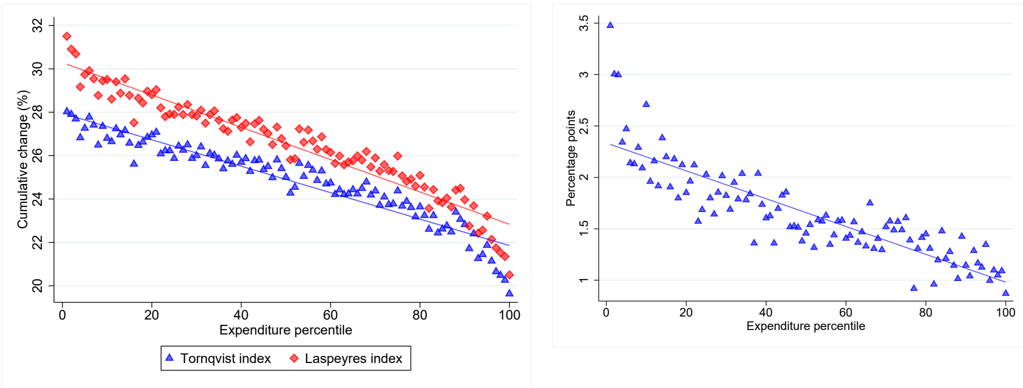
[21] If data on prices and shares are available in continuous time, then in principle it is possible to compute an exact cost-of-living index using the Divisia index (Divisia (1926)). With data available in discrete time it is necessary to approximate the cost-of-living index between periods of observation. Chaining a superlative index across periods of comparison approximates the Divisia index. However, in practice chain-drift bias often leads price indexes to perform poorly when chained at high frequency (see Ivancic et al. (2011), Diewert (2022)). In common with results in preceding sections, we chain indexes in this section across the calendar quarters 2021Q3, 2022Q3 and 2023Q3.

[22] A first-order approximation to $\log e^H_h(\mathbf{P}_t)$ around $\mathbf{P}_s$ implies $\log e^H_h(\mathbf{P}_t) - \log e^H_h(\mathbf{P}_s) \approx \sum_i s_{his}(\log P_{it} - \log P_{is})$ and a first-order approximation to $\log e^H_h(\mathbf{P}_s)$ around $\mathbf{P}_t$ implies $\log e^H_h(\mathbf{P}_t) - \log e^H(\mathbf{P}_s) \approx \sum_i s_{hit}(\log P_{it} - \log P_{is})$, where in both cases we have used Shephard's Lemma: $\frac{\partial \log e^H_h(\mathbf{P})}{\partial \log P_i} = s_{hi}$. Taking an arithmetic average leads to the second-order (Trapezoid) approximation: $\log \mathbb{P}^H_h(\mathbf{P}_s, \mathbf{P}_t) = \log e^H_h(\mathbf{P}_t) - \log e^H_h(\mathbf{P}_s) \approx \log \mathbb{P}^T_{h,(s,t)}$.

rises. This effect is quantitatively important but it does not unwind the large inflation gradient. For instance, the difference in cumulative inflation over 2021Q3-2023Q3 between households at the 80 and 20 expenditure percentiles based in the Törnqvist index is 3.3 percentage points (in comparison to 3.8 under the Laspeyres index).

Figure 5: *Inflation gradient accounting for substitution bias*

(a) Törnqvist index (b) Implied substitution bias



Notes: Authors' calculations using Kantar's Take Home panel (2021-2023). Panel (a) and (b) plot the relationship expenditure percentile with (in panel (a)) cumulative inflation over 2021Q3-2023Q3 computed using a Laspeyres and Törnqvist index and (in panel (b)) substitution bias in the Laspeyres price index. Substitution bias is measured as the difference in the Laspeyres and Törnqvist indexes. We allocate households to expenditure percentiles based on their equivalized spending in 2021. Cumulative inflation is computed using an annually chained index.

New product variety bias. A limitation of the Laspeyres index (shared by the Törnqvist index) is it only compares prices of products available in both comparison periods. This can give rise to a new product variety bias, since the indexes fail to account for welfare losses from product exits and gains from entry. In our data 7.1% of 2021Q3 aggregate spending is on products not available in 2022Q3 and 8.6% of 2022Q3 aggregate spending is on products that were not available in 2021Q3.[23] The analogous comparison between 2022Q3 and 2023Q3 results in 5.1% of 2022Q3 spending on exiting products and 6.5% of 2023Q3 spending on entering products. The Feenstra corrected CES price index (Feenstra (1994)) provides a convenient way of quantifying the importance of product entry and exit. If the household has CES preferences their cost-of-living index takes the form $\mathbb{P}^{CES}_{h,(s,t)} = \left[\prod_i \left(\frac{P_{it}}{P_{is}}\right)^{\phi_{hi,(s,t)}}\right] \times \mathcal{F}_{h,(s,t)}$, where $\phi_{hi,(s,t)}$ is a weight that depends on the household's period s and t spending shares over products available in both comparison periods.[24] $\mathcal{F}_{h,(s,t)} =$

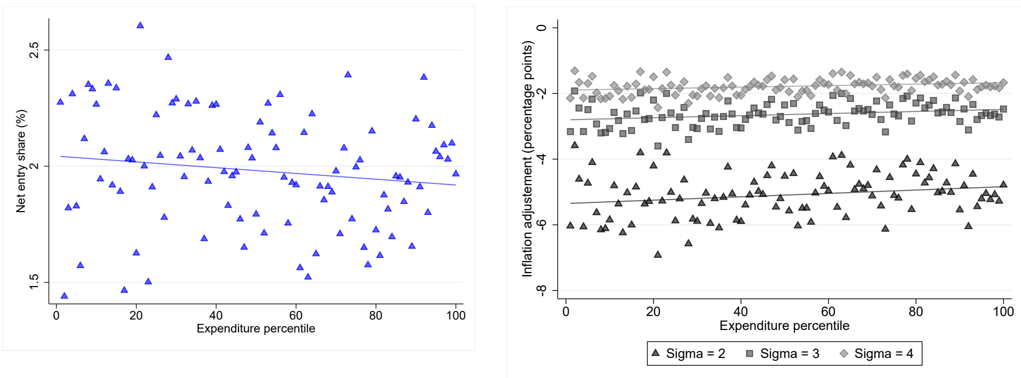
[23]We define a product as being unavailable in a quarter if no household in our entire sample purchases it.

[24]Specifically, $\phi_{hi,(s,t)} = \frac{(s_{hit}-s_{his})/(\log s_{hit}-\log s_{his})}{\sum_{i'}(s_{hi't}-s_{hi's})/(\log s_{hi't}-\log s_{hi's})}$.

$\left(\frac{1-s_{ht}^N}{1-s_{hs}^X}\right)^{\frac{1}{\sigma-1}}$ is the correction for the influence of entering and exiting products. It depends positively on the share of initial period spending the household allocates to exiting goods, s_{hs}^X (the cost-of-living grows if products the households consumes disappear) and negatively on the share of final period spending they allocate to entering goods, s_{ht}^N (as the consumption of new products contributes towards a lower cost-of-living). The sensitivity of the cost-of-living index to net product entry depends on the elasticity of substitution, $\sigma > 0$; when σ is low, exiting and entering products do not have close substitutes and therefore net entry has a relatively large effect of the cost-of-living. In Figure 6(a) we plot the difference in entry and exit shares, $s_{ht}^N - s_{hs}^X$ and in panel (b) we show the adjustment to inflation implied by net product entry by reporting the difference between the Feenstra corrected CES and baseline (no correction CES) index for a range of values of σ. In each case we report results for each percentile of the expenditure distribution. The figure shows that households allocated a higher share of their spending to entering than exiting products, which acts to lower cost-of-living increases. However, the magnitude of the effect is small (on average, even for a low substitution elasticity ($\sigma = 2$) net entry leads to a reduction of cost-of-living increases over 2021Q3-2021Q3 of slightly over 1 percentage point), and it is very similar across the expenditure distribution.

Figure 6: *Distributional effect of product entry and exit*

(a) Average net product entry (b) Inflation adjustment due for net entry



Notes: Authors' calculations using Kantar's Take Home panel (2021-2023). Panel (a) plots the average difference in entry and exit shares $s_{ht}^N - s_{hs}^X$ across expenditure percentiles, averaging over the periods 2021Q3-2022Q3 and 2022Q3-2023Q3. Panel (b) plots the difference between implied Feenstra corrected CES index and standard CES price index across expenditure percentiles. We show this difference for three different values of CES elasticity of substitution ("sigma"). We allocate households to expenditure percentiles based on their equivalized spending in 2021. Panel (b) is based annually chained indexes. We report the level of the unadjusted CES index in Appendix C.

Non-homotheticity bias. If preferences are non-homothetic, households will adjust their spending shares through an income effect arising from any changes in their standard of living, in addition to substitution effects arising from relative price changes. This imparts a non-homotheticity bias on superlative indexes. Recent evidence shows that non-homotheticity bias can be substantial when comparing standards of living over several decades (e.g., Comin et al. (2021), Baqaee et al. (2024), Jaravel and Lashkari (2024)). As the period 2021Q3-2023Q3 entailed sharp increases in the prices of fast-moving consumer goods as part of a broader "cost-of-living crisis", where prices in other sectors also rose and household incomes were stagnant, non-homotheticities in preferences may also play an important in driving cost-of-living changes over this time.

When preferences are non-homothetic, a second-order approximation to a cost-of-living index is given by the Hicksian share Törnqvist index, which in logs takes the form:[25]

$$\log \mathbb{P}^{HT}_{h,(s,t)}(u) = \frac{1}{2} \sum_i \left(\omega_{his}(u) + \omega_{hit}(u)\right) \left(\log P_{it} - \log P_{is}\right).$$

The difference with the standard Törnqvist index is that the spending shares are Hicksian shares evaluated at reference utility, $\omega_{hi\tau}(u)$, and are not directly observable. For concreteness, consider the direct comparison between the initial and final periods, 1 and T, with reference utility set to that realised in period 1, $u = u_{h1}$. This means that $\omega_{hi1}(u_{h1}) = s_{hi1}$, and the Hicksian share and standard Törnqvist indices differ only through the final shares ($\omega_{hiT}(u_{h1})$ in the former and s_{hiT} in the latter). If the household's utility is falling over time, then $s_{hiT} > \omega_{hiT}(u_{h1})$ for necessities and $s_{hiT} < \omega_{hiT}(u_{h1})$ for luxuries. Therefore, if product-level price growth and total expenditure elasticities are correlated the standard Törnqvist index may no longer approximate the cost-of-living index well.

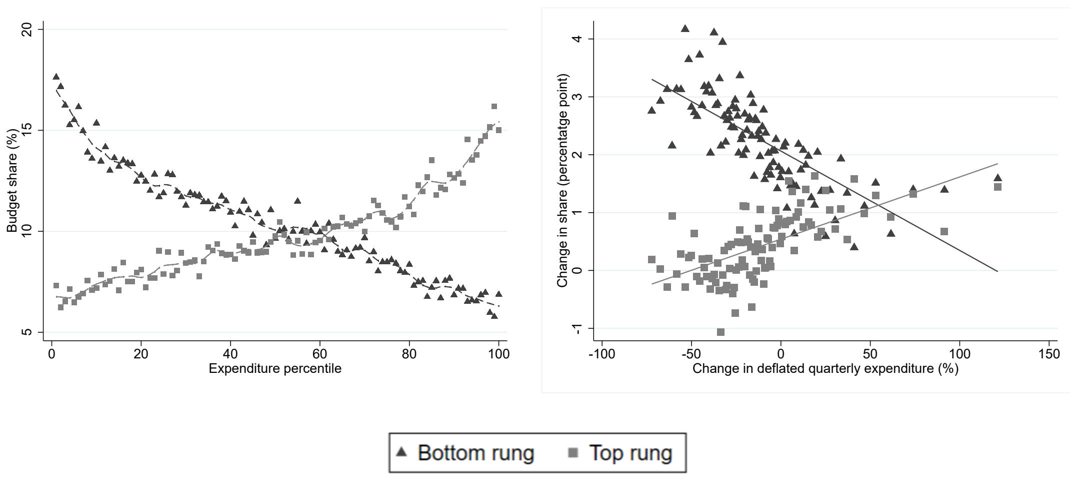
In Figure 7 we provide evidence that preference non-homotheticities were important in driving households' responses to the 2021-2023 cost-of-living crisis. Panel (a) shows that the share of their total quarterly expenditure households allocated to products on the bottom quality rung in 2021Q3 is strongly decreasing across the expenditure percentiles, while the top quality rung spending shares are strongly increasing across the percentiles. This suggests that low quality products are necessities and high quality products are luxuries.

[25] A first-order approximation to $\log e_h(\mathbf{P}_t, u)$ around $\mathbf{P}_s$ implies $\log e_h(\mathbf{P}_t, u) - \log e_h(\mathbf{P}_s, u) \approx \sum_i \omega_{his}(u)(\log P_{it} - \log P_{is})$ and a first-order approximation to $\log e_h(\mathbf{P}_s, u)$ around $\mathbf{P}_t$ implies $\log e_h(\mathbf{P}_t, u) - \log e_h(\mathbf{P}_s, u) \approx \sum_i \omega_{hit}(u)(\log P_{it} - \log P_{is})$, where in both cases we have used Shephard's Lemma: $\frac{\partial \log e_h(\mathbf{P},u)}{\partial \log P_i} = \omega_{hi}(u)$. Taking an arithmetic average leads to the second-order (Trapezoid) approximation: $\log \mathbb{P}_h(\mathbf{P}_s, \mathbf{P}_t, u) = \log e_h(\mathbf{P}_t, u) - \log e_h(\mathbf{P}_s, u) \approx \log \mathbb{P}^{HT}_{h,(s,t)}(u)$.

On average, households raised the share of their spending they allocated to both bottom and to top quality rung products between 2021Q3 and 2023Q3 (by 2.2 and 0.4 percentage points on average). The increase in top quality rung shares, which saw the lowest price growth, is consistent with substitution behaviour. The increase in bottom quality rung budget shares, despite these products becoming relatively more expensive, is suggestive of an income effect (trading down to products that are cheaper in absolute terms as living standards fall). Figure 7(b) shows that the *within*-household change in budget shares allocated to top and bottom quality rung products over 2021Q3-2023Q3 are systematically related to within-household changes in deflated expenditure. Those households with the largest falls in deflated expenditure switch most strongly towards low quality products and least strongly towards high quality products. This is consistent with households moving along the quality rung Engel curves in response to changes in their living standard, and suggests that the standard Törnqvist index will over-estimate the cost of maintaining living standards at their 2021Q3 level.

Figure 7: *Top and bottom quality rung Engel curves*

(a) Cross-sectional (b) Within-household



Notes: Authors' calculations using Kantar's Take Home panel (2021-2023). Panel (a) reports the average household spending share allocated to products belonging to the top and bottom quality rungs in 2021Q3, by expenditure percentile. The dashed lines are local polynomial-smoothed regressions. We allocate households to expenditure percentiles based on their 2021 equivalized spending. Panel (b) shows the average percentage point change in spending share allocated to products on the bottom and top quality rung between 2021Q1 to 2023Q3 for each percentile of the distribution of percent changes in deflated quarterly expenditure over this time. Expenditure changes are deflated using a Laspeyres price index.

To quantify the extent of non-homotheticity bias we use the approach developed in Jaravel and Lashkari (2024) (hereafter JL). We refer the reader to their paper for the full details, here sketching the details with a view to providing intuition. The method exploits the close connection between the standard-of-living (or quantity) index and the money-metric utility function.

The standard-of-living index between periods s and $t > s$ is defined as $\mathbb{Q}_h(u_{hs}, u_{ht}; \mathbf{P}) = \frac{e_h(\mathbf{P},u_{ht})}{e_h(\mathbf{P},u_{hs})}$; it measures the change in minimum expenditure needed to achieve realised utility when the price vector is fixed at $\mathbf{P}$. The money-metric utility function transforms the units of utility from arbitrary "utils" to money, and is defined by $\mathbb{u} = e_h(\mathbf{P}_b, u)$, i.e., the expenditure needed at some base price vector $\mathbf{P}_b$ to achieve u utils. If a household faces the price vector $\mathbf{P}_t$ and has nominal expenditure x, which enables them to achieve $u_{ht} = v_h(\mathbf{P}_t, x)$ utils, these utils can be converted into monetary terms according to $\mathbb{u}_{ht} = e_h(\mathbf{P}_b, u_{ht}) = e_h(\mathbf{P}_b, v_h(\mathbf{P}_t, x))$. Similar logic applies to the expenditure function; at prices $\mathbf{P}_t$ the expenditure necessary to achieve u utils (or its money-metric equivalent, $\mathbb{u} = e_h(\mathbf{P}_b, u)$) is $e_h(\mathbf{P}_t, u) = e_h(\mathbf{P}_t, v_h(\mathbf{P}_b, \mathbb{u}))$. The JL algorithm constructs a sequence of money-metric utilities $(\mathbb{u}_{h1}, \ldots, \mathbb{u}_{hT})$, which in turn are a base price b standard-of-living index.

Consider the change in log expenditure between periods t and $t-1$, which can be written:

$$\begin{aligned}\log x_{ht} - \log x_{ht-1} = [\log e_h(\mathbf{P}_t, v_h(\mathbf{P}_b, \mathbb{u}_{ht-1})) - \log e_h(\mathbf{P}_{t-1}, v_h(\mathbf{P}_b, \mathbb{u}_{ht-1}))] +\\ [\log e_h(\mathbf{P}_t, v_h(\mathbf{P}_b, \mathbb{u}_{ht})) - \log e_h(\mathbf{P}_t, v_h(\mathbf{P}_b, \mathbb{u}_{ht-1}))] \quad (5.1)\end{aligned}$$

This decomposes expenditure growth into two components. The first reflects price changes at period $t-1$ utility (and is a log cost-of-living index evaluated at period $t-1$ utility). The second reflects the change in the household's utility over time at period t prices (and is a log standard-of-living index evaluated at period t prices).

A first-order approximation to the cost-of-living index is given by $\log \mathbb{P}^{GL}_{h,(t-1,t)} = \sum_i s_{hit-1}(\log P_{it} - \log P_{it-1})$, i.e., a geometric-Laspeyres price index, and a first-order approximation to the standard-of-living index is given by $\frac{\partial \log e_h(\mathbf{P}_t, v_h(\mathbf{P}_b, \mathbb{u}_{ht-1}))}{\partial \log \mathbb{u}}(\log \mathbb{u}_{ht} - \log \mathbb{u}_{ht-1})$. While the first approximation depends on data, the second depends on unobservables (namely utility and the expenditure function). However, this can be overcome by setting base prices for the money-metric utility function at either period 1 or period T prices. For concreteness, consider initial period base prices. This means that initial period money-metric utility is observable and is given by $\mathbb{u}_{h1} = x_{h1}$, and the derivative of the expenditure function (the non-homotheticity adjustment) can by replaced by: $\frac{\partial \log e_h(\mathbf{P}_t, v_h(\mathbf{P}_1, \mathbb{u}_{ht-1}))}{\partial \log \mathbb{u}} = \frac{\partial \log \mathbb{P}_h(\mathbf{P}_1, \mathbf{P}_t, \mathbb{u}_{ht-1})}{\partial \log \mathbb{u}} + 1$.[26]

[26] The first claim follows from the fact that expenditure needed to achieve money-metric utility $\mathbb{u}_{h1}$ in the initial period is $e_h(\mathbf{P}_1, v(\mathbf{P}_b, \mathbb{u}_{h1}))$ along with choice of base prices $\mathbf{P}_b = \mathbf{P}_1$. The second claim follows from the definition of the cost-of-living index: $\log \mathbb{P}_h(\mathbf{P}_1, \mathbf{P}_t, \mathbb{u}_{ht-1}) = \log e_h(\mathbf{P}_t, v(\mathbf{P}_b, \mathbb{u}_{ht-1})) - \log e_h(\mathbf{P}_1, v(\mathbf{P}_b, \mathbb{u}_{ht-1}))$ along with choice of base prices $\mathbf{P}_b = \mathbf{P}_1$.

Under the assumption that the relationship between money-metric utility and the cost-of-living index is common across households, $\frac{\partial \log \mathbb{P}_h(\mathbf{P}_1,\mathbf{P}_t,\mathbb{u}_{ht-1})}{\partial \log \mathbb{u}} \equiv \Lambda_t(\mathbb{u}_{ht-1})$, $\Lambda_2(\mathbb{u}_{h1})$ can be recovered by a cross-sectional non-parametric (series) regression of household-level (geometric-Laspeyres) price indexes between periods 1 and 2 and $\mathbb{u}_{h1}$. The sequence of money-metric utilities can then be constructed recursively. For instance, the period 2 money-metric utility can recovered from $\log \hat{\mathbb{u}}_{h2} = \log \mathbb{u}_{h1} + \frac{1}{1+\hat{\Lambda}_2(\mathbb{u}_{h1})}\left(\log x_{h2} - \log x_{h1} - \log \mathbb{P}^{GL}_{1,2}\right)$. Finally, the cost-of-living index between period 1 and t, evaluated at period u_{ht} utility, $\mathbb{P}_h(\mathbf{P}_1, \mathbf{P}_t; u_{ht})$, can be recovered from $\log \mathbb{P}_h(\mathbf{P}_1, \mathbf{P}_t; u_{ht}) = \log x_{ht} - \log \hat{\mathbb{u}}_{ht}$.[27]

This algorithm provides a non-parametric approximation to a non-homothetic cost-of-living index. The key assumptions underlying it are that a) product-level Engel curves are common across households (which means that the function $\Lambda_t(.)$ is common across households) and b) conditional on control variables included in the regression of price indexes on money-metric utility, differences in inflation rates across households with different levels of money-metric utility (which arise from differences in their chosen consumption baskets) are driven by preference non-homotheticities rather than preference heterogeneity. When implementing the algorithm we control for the demographic composition of households (number of children aged below 14 and aged between 14 and 18, and number of adults).

As described, this approach builds a cost-of-living comparison between periods 1 and T with utility held at the final period realised value, and adjusts a geometric-Laspeyres price index to allow for non-homotheticities in preferences. In practice we implement a modified version of this approach (also outlined in JL). It entails instead constructing a period 1 and T cost-of-living comparison with utility fixed at its *initial* period level. In our context we find it natural to document how the cost of achieving a household's initial utility level varied over the 2021Q3-2023Q3 inflation. We also adjust the Törnqvist price index for non-homotheticities. We find this conceptually more attractive than using the less flexible (non-superlative) geometric-Laspeyres index.[28]

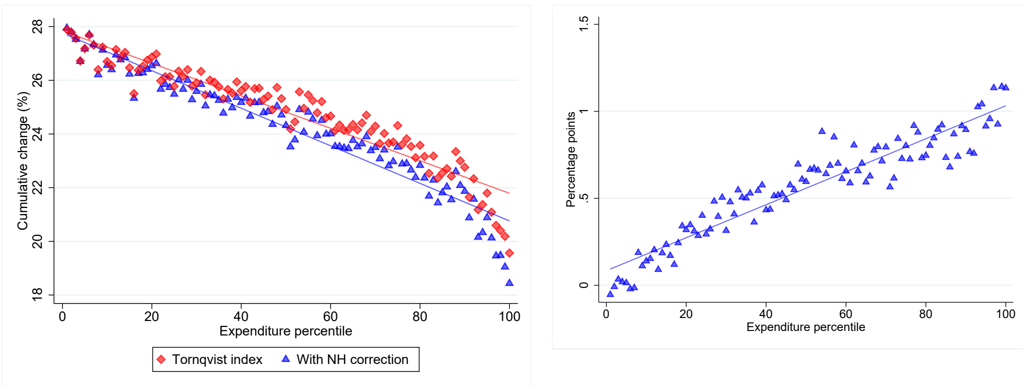
[27] To see this note that $\log x_{ht} = e_h(\mathbf{P}_t, u_{ht})$ and $\log \mathbb{u}_{ht} = \log e_h(\mathbf{P}_1, u_{ht})$.

[28] Constructing an initial period cost-of-living index entails choosing final period base prices for money-metric utility and running the recursive algorithm in reverse, i.e., starting from $\mathbb{u}_{hT} = x_{hT}$ and moving backwards in time. In Appendix A we motivate the adjustment to the Törnqvist index by combining the approximation of equation (5.1) described in the text, with an analogous approximation, applied to a decomposition of $\log x_{ht} - \log x_{ht-1}$, into a period t base utility cost-of-living index and a period $t-1$ base price standard-of-living index. Prior to implementing the algorithm we average across households within each percentile of the 2021 equivalized expenditure distribution, to reduce noise from mean-reversion in the household-level data. We implement the reverse first-order version of the JL algorithm, deflating nominal expenditure growth with a Törnqvist index. Consistent with the price indexes throughout the paper, we implement the algorithm comparing the calendar quarters 2021Q3, 2022Q3 and 2023Q3.

In Figure 8(a) we report the inflation gradient under the standard Törnqvist index and the version corrected for non-homothetic preferences. In panel (b) we show the implied non-homotheticity bias in the non-corrected Törnqvist index. The figure shows that the Törnqvist index over-estimates cost-of-living increases, with the degree of bias increasing across the expenditure distribution. Therefore the Törnqvist index understates the extent of substitution effects, especially for the better-off, as it contaminated by these households switching to lower quality, high price growth, products due to the income effect associated with their living standard falling. Existing evidence has shown that it is important to account for preference non-homotheticities when measuring changes in living standards over several decades. Our results show that preference non-homotheticities can play an important role over relatively short time horizons when inflation rates are high.

Figure 8: *Inflation gradient accounting for non-homotheticity bias*

(a) Non-homothetic index (b) Implied non-homotheticity bias



Notes: Authors' calculations using Kantar's Take Home panel (2021-2023). Figure plots the relationship between the percentile of the expenditure distribution a household belongs to with (in panel (a)) cumulative inflation over 2021Q3-2023Q3 computed using a standard Törnqvist and one with a non-homotheticity correction and (in panel (b)) non-homotheticity bias in the Törnqvist price index. Non-homotheticity bias is measured as the difference in the standard and corrected Törnqvist indexes. We allocate households to expenditure percentiles based on their 2021 equivalized expenditure. Cumulative inflation is computed using an annually chained index.

Overall, after accounting for substitution effects and non-homotheticities, households in the top decile of the expenditure distribution experienced cost-of-living rises that were 7.2 percentage points lower than those in the bottom decile.

6 Cost pass-through

The 2021-2023 inflationary spike was driven by a set of inter-related supply shocks, which contributed to rises in the cost of several factors of production, including energy, transportation and wholesale commodity prices, as well as to labour shortages. It is possible that these fed through into bigger (proportional) marginal cost rises for lower quality products. Alternatively, or additionally, the extent to which firms passed-through marginal cost rises may have generated differential proportional price increases across the quality ladder.

In much of the macro and trade literature the benchmark pricing model entails monopolistic firms that produce at constant marginal cost and face CES demand (Dixit and Stiglitz (1977)). In this case a firm sets its price at a constant multiplicative mark-up over cost, and in level terms cost shocks are over-shifted to prices, while in proportional terms cost pass-through is one.[29] Therefore this benchmark model rules out differential proportional pass-through as a source of varying price increases across the quality ladder. However, CES demands impose very strong restrictions on demand curvature and hence cost pass-through (see Mrázová and Neary (2017)), while empirical evidence points towards pass-through of cost shocks that is less than proportional (e.g., Gopinath and Itskhoki (2010), De Loecker et al. (2016), Miravete et al. (2018)).

As an example of how differential proportional cost pass-through may give rise to cheapflation, consider a high (h) and low (l) quality product, produced at marginal cost c_j, where $j = \{l, h\}$ and denote their price $P_j = c_j + \psi_j$, where ψ_j is the additive price-cost margin (and where $P_h > P_l$). Suppose that $\psi_h > \psi_l$ (as suggested by recent evidence by the UK Competition and Markets Authority (2023)). If pass-through of cost shocks in levels is the same across products ($\frac{dP_j}{dc_j} = \rho$) [30] then in proportional terms pass-through of the cost shock will be higher for the low quality product if $\frac{\psi_h}{\psi_l} > \frac{c_h}{c_l}$ (i.e., if, in proportional terms, price-cost margins rise more quickly across the quality ladder than marginal costs).

In reality pass-through (in levels or proportionately) may vary across products, depending on a range of factors, including a product's relative demand and supply elasticity, demand curvature, the competitiveness of conduct in the market and multiproduct firm portfolio effects (see Hamilton (2009); Weyl and Fabinger (2013)). In this section we study an earlier, more specific, cost shock and assess whether there is any evidence that it generated differential inflation across cheaper and more expensive products.

[29] A monopolist with constant marginal cost c facing CES demand $q = AP^{-\sigma}$, where $\sigma > 1$ will set price $P = \frac{\sigma}{\sigma-1}c$ with level pass-through $\frac{dP}{dc} = \frac{\sigma}{\sigma-1}$ and proportional pass-through $\frac{d\log P}{d\log c} = 1$.

[30] The family of demand functions due to Bulow and Pfleiderer (1983) give rise to constant level pass-through under monopolistic pricing. Sangani (2023) provide recent evidence of complete level pass-through in grocery markets. He also writes down a model that rationalises retailers setting fixed price-cost margins on the basis that they face uncertain demand and overhead expenses, and wish to mitigate the risk of variable profits failing to cover overheads (rather than through the curvature properties of demand functions).

6.1 The Brexit shock

We focus on the unusually rapid and persistent depreciation of the sterling-euro exchange rate around the time of the UK's referendum on continuing membership of the European Union on June 23^{rd} 2016 (the "Brexit" referendum). This led to a marginal cost shock for products imported from EU countries. We focus on fresh produce, and compare the proportional pass-through of the exchange rate shock to prices between initially low and high priced imports from the EU. We interpret higher proportional pass-through of the exchange rate shock for lower priced products as evidence in favour of higher proportional pass-through of a marginal cost shock for these products.

If the marginal cost of EU imports comprises only the cost of importing the produce, then exchange rate pass-through to prices will coincide with marginal cost pass-through to prices. In reality there is likely to also be domestic inputs (e.g., domestic transport and storage cost), that are complementary in production to the imported produce, which will create a wedge between exchange rate and marginal cost movements (Goldberg and Hellerstein (2013)). For instance, suppose a composite domestic input factor, priced at w_d, is combined in fixed proportions with imported produce, priced w_f (and the input quantities are normalised such that they are combined one-of-one). In this case the effect of a proportional increase in import costs (due to an exchange rate depreciation) on marginal cost, mc, will be $\frac{d \log mc}{d \log w_f} = \frac{w_f}{w_f + w_d}$. Therefore, if the domestic cost share varies across products, it will result in differential proportional pass-through of exchange rate movements to marginal costs (higher domestic cost shares will result in lower pass-through). As products across the quality ladder are sold by the same retailers, we think it most likely that high quality products will have lower domestic cost shares and therefore higher proportional exchange rate pass-through to marginal cost. This would then act to reinforce the interpretation that higher proportional exchange rate pass-through to the price of cheap products reflects higher proportional marginal cost pass-through to prices.

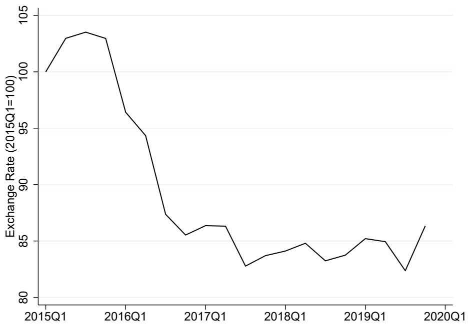
In Figure 9(a) we report the Pound-Euro exchange rate over the period 2015-2019.[31] Over 2016Q1 to 2016Q4, the relative value of the pound fell by 20.1%. The pound then remained weak and stable relative to the euro for several years. This depreciation was driven by opinion polls, which had initially shown a clear lead for continued EU membership, tightening during the referendum campaign (which

[31] Not all EU members are part of the Eurozone, however of those that are not two (out of eight) maintain currency pegs with the Euro and all saw similar currency appreciations relative to the pound over this period.

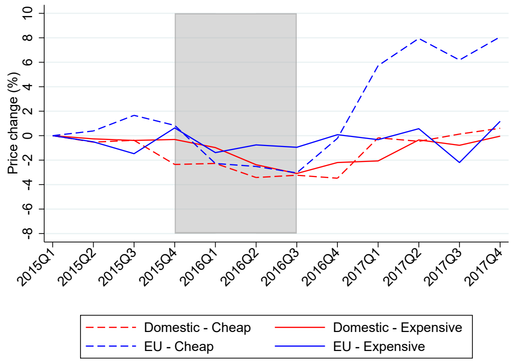
officially began on April 15^{th}, i.e., 2016Q2), as well as the surprise caused by the referendum result itself, which was also accompanied by sharp falls in the stock market (Breinlich et al. (2018)).

Figure 9: *Impact of Brexit exchange rate shock*

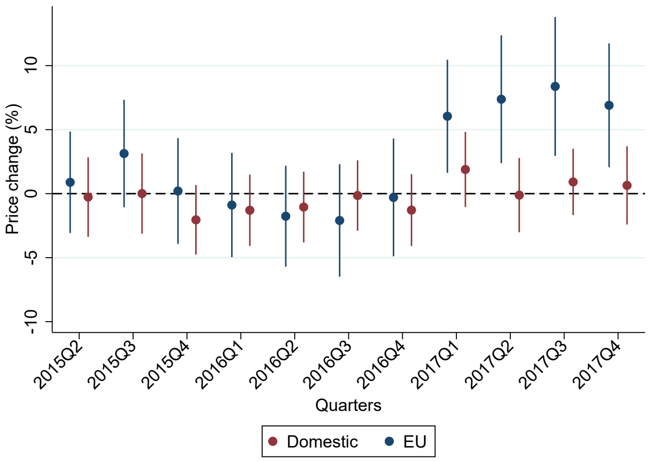
(a) GBP-EUR exchange rate



(b) Price changes



(c) Difference-in-difference results



Notes: Authors' calculations using Kantar's Take Home panel (2015-2017) and exchange rate data from the Office for National Statistics. Panel (a) plots the Pound-Euro exchange rate. Panel (b) plots the evolution of average prices for domestic and EU produce, split by whether they are in the top ("expensive") or bottom ("cheap") half of the average price distribution over 2015. The shaded area shows highlights the period of Pound depreciation. Panel (c) report estimates and 95% confidence bands from the equation (6.1), which is estimated separately on domestic and EU products.

In panel (b) of Figure 9 we show how the retail prices of fresh products with identifiable countries of origin evolved over this period over 2015Q1-2017Q4. To do this we focus on the set of domestic and EU-imported products available throughout this time period[32] and for both domestic and EU imports, we split products by whether they are relatively "cheap" or "expensive" based on whether, conditional on product category, their average price over 2015 lies in the bottom or top half

[32]Comprising 53.8% of total expenditure for categories with country of origin information.

of the average price distribution (across both domestic and EU produce). The graph shows the evolution of average price for cheap and expensive domestic and EU goods (weighted by their 2015Q1 spending share). While the average price of expensive EU products is relatively constant throughout this period (and evolves roughly similarly to that of expensive domestic products), the average price of cheap EU products rises by close to 10% in the 4 quarter period that followed the sterling depreciation.[33]

In panel (c) we report estimates from two regressions (one for domestic and one EU products) that compare the evolution of the price of cheap relative to expensive products. Specifically, separately for the set of domestic and EU products, we estimate the equation:

$$\tilde{P}_{it} = \beta_0 + \sum_t \beta_{1t} D_t + \sum_t \beta_{2t} D_t \times \text{Cheap}_i + \alpha_i + \varepsilon_{it} \tag{6.1}$$

where $\tilde{P}_{it}$ is product i's year-quarter price (relative to 2015Q1), D_t are year-quarter dummies, Cheap_i is an indicator variable for whether a product is in the lower half of the 2015 average price distribution, α_i are product fixed effects, and ε_{it} is a product-time deviation. Panel (c) reports the estimate of β_2. It shows that while prices rose by similar proportional amounts for cheap and expensive domestic products, proportional price growth for cheap EU products outstripped that for expensive EU products.[34]

In sum, Figure 9 shows that the Brexit-referendum-induced exchange rate depreciation led to higher price growth among initially low priced products. As long as the exchange rate shock did not lead to a proportional marginal cost shock that is significantly lower for expensive than for cheaper products (which would happen under the, unlikely in our view, scenario of more expensive fresh products having higher domestic cost shares), this points toward higher proportional pass-through of marginal cost shocks for lower quality products. The cost shocks that drove the 2021-2023 inflation are more complicated and would have impacted all products. However, the evidence of how firms adjusted prices as a result of the Brexit shock, coupled with the striking pattern of differential price changes across the quality ladder during the 2021-2023 inflation, point towards unequal proportional pass-

[33] Sluggish adjustment to exchange rates shocks is a common empirical finding (e.g., Nakamura and Zerom (2010)).

[34] A Wald test, based on estimating equation (6.1) jointly, with coefficients fully interacted with domestic and EU dummy, suggests the coefficients on year-quarter dummy variables are significantly different at 95% level between domestic and EU produce from 2017Q2 to 2017Q4.

through of marginal cost changes being an important driver of inflation inequality over this period.

7 Conclusion

We show that the 2021-2023 cost-of-living crisis entailed a steep gradient in household-level inflation rates. The pattern is unprecedented in the UK in recent times and the magnitude of the gradient is substantially larger than that documented in other countries during times of moderate inflation. The gradient is driven by differences in consumption baskets across products within narrowly defined product categories, underlining the importance of using granular data to study inflation inequality (Jaravel (2021)) and in answer to calls for income-group specific inflation indices (Klick and Stockburger (2021, 2024)). The inflation inequality is driven by a novel pattern of cheapflation; higher proportional prices rises for lower quality products, that comprise a relatively high share of less well-off households' consumption baskets. Households responded both to relative price changes and to changes in their living standards, which led them to switch to lower quality, high price growth products. Therefore both substitution effects and preference non-homotheticities played an important role in driving expenditure switching over this period. After accounting for them we find the gradient in the cost of maintaining living standards at their pre-inflationary shock level exhibits a similar significant gradient. The causes of differential price growth across the quality ladder over 2021-2023 are likely to be complex and an important area for future research. Using the previous exchange rate shock around the Brexit vote, we show exchange rate changes were passed through more in proportional terms to the prices of cheaper products, indicating that higher proportional pass-through of marginal cost shocks is likely to be an important part of the story.

Recent evidence suggests that the immediate post COVID-19 pandemic period bore witness to cheapflation in a number of countries (Cavallo and Kryvtsov (2024)). This raises the possibility that cheapflation has fueled inflation and cost-of-living inequality in many countries. It also remains to be seen whether this pattern will persist in periods of more general price stability. Both topics are important avenues for future research.

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Appendix:
For online publication

Cheapflation and the rise of inflation inequality

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August, 2024

A Non-homotheticity correction

In this section we provide further details of the non-homotheticity correction that we implement in Section 5 and that is developed in Jaravel and Lashkari (2024) (we refer the reader to this paper for full details).

Definitions. Denote household h's indirect utility function by $v_h(\mathbf{P}, x)$ and its inverse, the expenditure function, by $e_h(\mathbf{P}, u) = v_h^{-1}(\mathbf{P}, .)$, where $\mathbf{P}$ is a price vector, x denotes total expenditure and u the corresponding utility level. Define the money-metric utility function as $\mathbb{u} = e_h(\mathbf{P}_b, u)$. If a household faces the price vector $\mathbf{P}_t$ and has nominal expenditure x, their money-metric utility is $\mathbb{u}_{ht} = e_h(\mathbf{P}_b, u_{ht}) = e_h(\mathbf{P}_b, v_h(\mathbf{P}_t, x))$. At prices $\mathbf{P}_t$ the expenditure necessary to achieve $\mathbb{u}$ is $e_h(\mathbf{P}_t, u) = e_h(\mathbf{P}_t, v_h(\mathbf{P}_b, \mathbb{u}))$.

First-order algorithm. Log nominal expenditure growth can be decomposed according to:

$$\log x_{ht} - \log x_{ht-1} = [\log e_h(\mathbf{P}_t, v_h(\mathbf{P}_b, \mathbb{u}_{ht-1})) - \log e_h(\mathbf{P}_{t-1}, v_h(\mathbf{P}_b, \mathbb{u}_{ht-1}))] + \\ [\log e_h(\mathbf{P}_t, v_h(\mathbf{P}_b, \mathbb{u}_{ht})) - \log e_h(\mathbf{P}_t, v_h(\mathbf{P}_b, \mathbb{u}_{ht-1}))]$$

Applying a first-order approximation to the expressions in square brackets (and Shephard's Lemma), this becomes:

$$\log x_{ht} - \log x_{ht-1} \approx \sum_i s_{hit-1}(\log P_{it} - \log P_{it-1}) + \\ \frac{\partial \log e_h(\mathbf{P}_t, v_h(\mathbf{P}_b, \mathbb{u}_{ht-1}))}{\partial \log \mathbb{u}}(\log \mathbb{u}_{ht} - \log \mathbb{u}_{ht-1}). \tag{A.1}$$

Note that

$$\log \mathbb{P}_h(\mathbf{P}_1, \mathbf{P}_t, \mathbb{u}_{ht-1}) = \log e_h(\mathbf{P}_t, v(\mathbf{P}_b, \mathbb{u}_{ht-1})) - \log e_h(\mathbf{P}_1, v(\mathbf{P}_b, \mathbb{u}_{ht-1}))$$

If we choose $b = 1$, then $\frac{\partial \log \mathbb{P}_h(\mathbf{P}_1, \mathbf{P}_t, \mathbb{u}_{ht-1})}{\partial \log \mathbb{u}} \equiv \Lambda_{ht}(\mathbb{u}_{ht-1}) = \frac{\partial \log e_h(\mathbf{P}_t, v(\mathbf{P}_1, \mathbb{u}_{ht-1}))}{\partial \log \mathbb{u}} - 1$ allows us to re-write equation (A.1) as:

$$\log x_{ht} - \log x_{ht-1} \approx \log \mathbb{P}^{GL}_{h,(t-1,t)} + (1 + \Lambda_{ht}(\mathbb{u}_{ht-1}))(\log \mathbb{u}_{ht} - \log \mathbb{u}_{ht-1}),$$

where $\log \mathbb{P}^{GL}_{h,(t-1,t)} \equiv \sum_i s_{hit-1}(\log P_{it} - \log P_{it-1})$. Finally restricting $\Lambda_{ht}(.) = \Lambda_t(.)$ for all h, this implies the recursive "first-order" sequence:

$$\begin{aligned}
\log \mathbb{u}_{h1} &= \log x_{h1} \\
\log \mathbb{u}_{h2} &= \log \mathbb{u}_{h1} + \frac{1}{1+\Lambda_2(\mathbb{u}_{h1})}\left(\log x_{h2} - \log x_{h1} - \log \mathbb{P}^{GL}_{h,(1,2)}\right) \\
&\dots \\
\log \mathbb{u}_{hT} &= \log \mathbb{u}_{hT-1} + \frac{1}{1+\Lambda_T(\mathbb{u}_{hT-1})}\left(\log x_{hT} - \log x_{hT-1} - \log \mathbb{P}^{GL}_{h,(T,T-1)}\right)
\end{aligned}$$

where at each stage one estimates $\log \mathbb{P}^{GL}_{h,(t-1,t)} = g_t(\log \mathbb{u}_{ht-1}) + \epsilon_h$ and constructs: $\Lambda_t(\mathbb{u}_{ht-1}) = \sum_{\tau \le t} g'_\tau(\log \mathbb{u}_{h1-t})$. The cost-of-living index comparing periods 1 and T is then obtained from:

$$\log \mathbb{P}_h(\mathbf{P}_1, \mathbf{P}_T; u_{hT}) = \log x_{hT} - \log \mathbb{u}_{hT}$$

Reverse first-order algorithm. The algorithm as described constructs a cost-of-living index evaluated at final period realized utility. The cost-of-living index evaluated at initial period utility can be constructed by choosing $b = T$ and running the algorithm in reverse (stating at period T). In this case the money-metric utility is evaluated at final base prices and the relevant cost-of-living index is obtained from $\log \mathbb{P}_h(\mathbf{P}_1, \mathbf{P}_T; u_{h1}) = \log \mathbb{u}_{h1} - \log x_{h1}$.

Second-order algorithm. A slightly modified approximation to equation (A.1) approximates the second term in parenthesis around $\mathbb{u}_{ht}$, giving:

$$\begin{aligned}
\log x_{ht} - \log x_{ht-1} \approx \sum_i & s_{hit-1}(\log P_{it} - \log P_{it-1}) + \\
& \frac{\partial \log e_h(\mathbf{P}_t, v_h(\mathbf{P}_b, \mathbb{u}_{ht}))}{\partial \log \mathbb{u}}(\log \mathbb{u}_{ht} - \log \mathbb{u}_{ht-1}). \qquad \text{(A.2)}
\end{aligned}$$

We can also decompose nominal expenditure growth according to:

$$\begin{aligned}
\log x_{ht} - \log x_{ht-1} = & [\log e_h(\mathbf{P}_t, v_h(\mathbf{P}_b, \mathbb{u}_{ht})) - \log e_h(\mathbf{P}_{t-1}, v_h(\mathbf{P}_b, \mathbb{u}_{ht}))] - \\
& [\log e_h(\mathbf{P}_{t-1}, v_h(\mathbf{P}_b, \mathbb{u}_{ht})) - \log e_h(\mathbf{P}_{t-1}, v_h(\mathbf{P}_b, \mathbb{u}_{ht-1}))]
\end{aligned}$$

Applying a first-order approximation to the expressions in square brackets, this becomes:

$$\log x_{ht} - \log x_{ht-1} = \sum_i s_{hit}(\log P_{it} - \log P_{it-1}) + \\ \frac{\partial \log e_h(\mathbf{P}_{t-1}, v_h(\mathbf{P}_b, \mathbb{u}_{ht-1}))}{\partial \log \mathbb{u}}(\log \mathbb{u}_{ht} - \log \mathbb{u}_{ht-1}). \quad \text{(A.3)}$$

If we choose $b=1$ use $\frac{\partial \log \mathbb{P}_h(\mathbf{P}_1,\mathbf{P}_t,\mathbb{u}_{ht})}{\partial \log \mathbb{u}} \equiv \Lambda_{ht}(\mathbb{u}_{ht}) = \frac{\partial \log e(\mathbf{P}_{ht}, v(\mathbf{P}_1,\mathbb{u}_{ht}))}{\partial \log \mathbb{u}} - 1$. Restricting $\Lambda_{ht}(.) = \Lambda_t(.)$ for all h and taking the arithmetic mean of equations (A.2) and (A.3) gives:

$$\log x_{ht} - \log x_{ht-1} \approx \log \mathbb{P}^T_{h,(t-1,t)} + \\ \left(1 + \frac{1}{2}\left(\Lambda_{t-1}(\mathbb{u}_{ht-1}) + \Lambda_t(\mathbb{u}_{ht})\right)\right)(\log \mathbb{u}_{ht} - \log \mathbb{u}_{ht-1}).$$

where $\log \mathbb{P}^T_{h,(t-1,t)} = \frac{1}{2}\sum_i (s_{hit-1} + s_{hit})(\log P_{it} - \log P_{it-1})$.

Compared with the first-order algorithm, this replaces the first-order price index $\log \mathbb{P}^{GL}_{h,(t,t-1)}$ with the second-order price index $\log \mathbb{P}^T_{h,(t,t-1)}$ and it replaces the first-order approximation to the derivative of the true log cost-of-living index, $\Lambda_t(\mathbb{u}_{ht-1})$, with a second-order approximation, $0.5 \times (\Lambda_{t-1}(\mathbb{u}_{ht-1}) + \Lambda_t(\mathbb{u}_{ht}))$, which comes with the additional computation cost of solving a fixed point problem for each (h,t).

B Alternative quality rungs

Here we document an alternative procedure for assigning products to quality rungs that controls for the impact of nonlinear pricing across different sizes of the same brand.

For each product category, over 2021Q3-2023Q3, we estimate the expenditure-weighted regression:

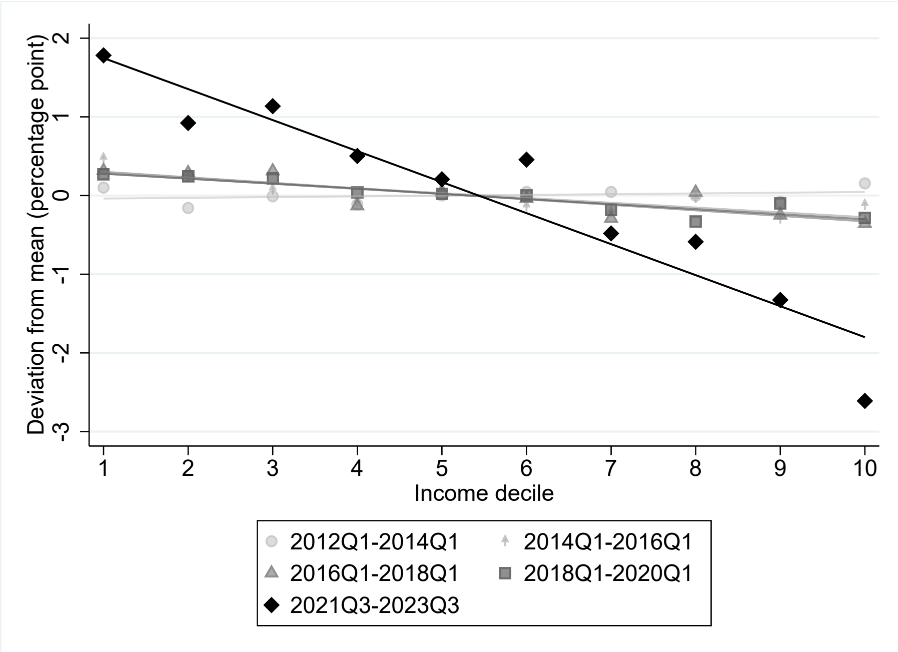
$$P_{it} = \xi_{b(i)} + \tau_{c(i)t} + \sum_y \sum_{l=1}^{3} \alpha^{(l)}_{c(i)y} \mathbb{1}\{t \in y\} \times \text{size}^{(l)}_i + \epsilon_{it}, \quad \text{(B.1)}$$

where P_{it} is the average quarterly price (per unit volume) paid for product i in quarter t, $\tau_{c(j)t}$ are year-quarter fixed effects, $\xi_{b(j)}$ are brand effects and the α terms represent a third-order polynomial in demeaned pack size with coefficients that vary across years (indexed by y).

To assign products across quality rungs, we first take products available in the four quarters 2021Q3-2022Q2 and compute their adjusted prices in each of these quarters netting off the size polynomial; $t \leq 4$, $\tilde{P}_{it} = P_{it} - \sum_{l=l} \alpha^{(l)}_{m(i)y} \mathbb{1}\{t \in y\} \times \text{size}^{(l)}_j$. We then average across quarters and use the expenditure-weighted distribution of adjusted prices within each product category to define decile boundaries. For all products including those enter after 2021Q2, we use equation (B.1) to impute the adjusted price at the average of first 4 quarters (i.e., we assign product j the adjusted price given by $\tilde{P}_{it} = \xi_{b(i)} + \frac{1}{4}\sum_{t=1}^{4} \tau_{c(i)t}$), and use this to assign the product to quality rungs. We measure the quality ladder analogously for the four earlier comparison periods.

C Results Appendix

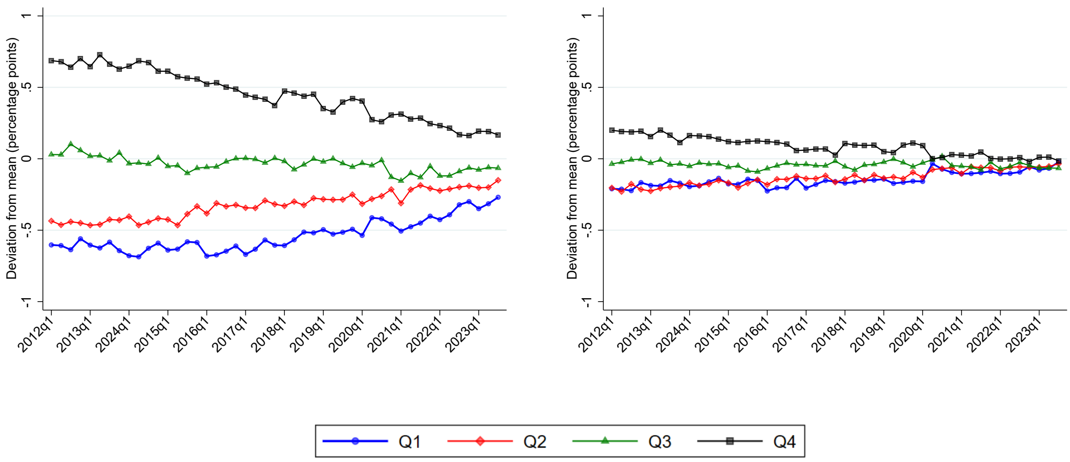
Figure 10: *Inflation inequality; across income deciles*



Notes: Authors' calculations using Kantar's Take Home panel (2012-2023). Figure plots the relationship between 9^{th} quarter cumulative inflation and income deciles based on a household's equivalized income computed based on the band midpoint of banded household income in the initial calendar year of the relevant nine quarter period. Cumulative inflation is computed using an annually chained Laspeyres index. The corresponding figure in the paper (based income measured as expenditure percentiles) is Figure 1(c).

Figure 11: *Household-level price dispersion*

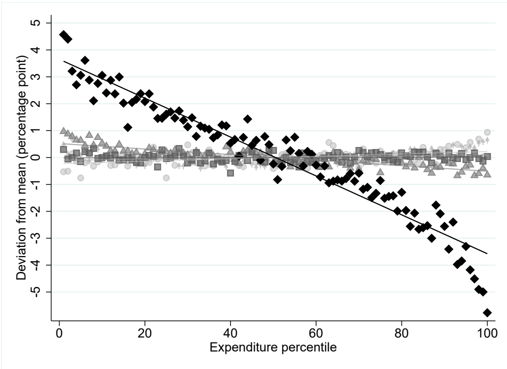
(a) AH index (product) (b) AH index (product-retailer)



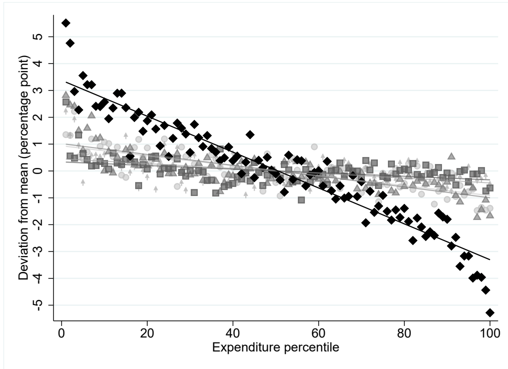
Notes: Authors' calculations using Kantar's Take Home panel (2012-2023). Figure show within income quartile average (across households) AH index at quarterly frequency over 2012Q1-2023Q3. In panel (a) products are based on brand-size and in panel (b) they are based on brand-size and retailer. Income quartiles are based on households annual equivalized income computed based on the band midpoint of banded household income. The corresponding figure in the paper (based income measured as expenditure percentiles) is Figure 2(a) and (b).

Figure 12: *Inflation inequality with alternative price indices*

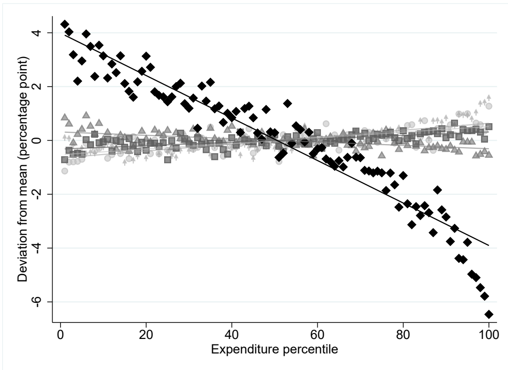
(a) Direct comparison Laspeyres



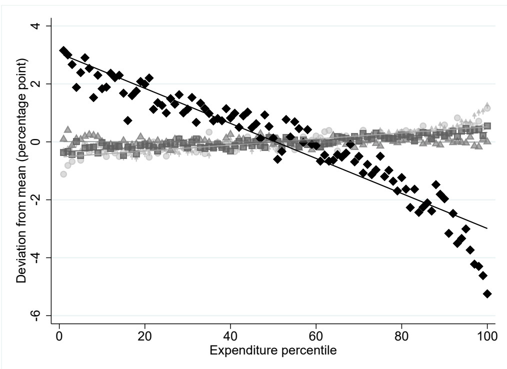
(b) Quarter-to-quarter chained Laspeyres



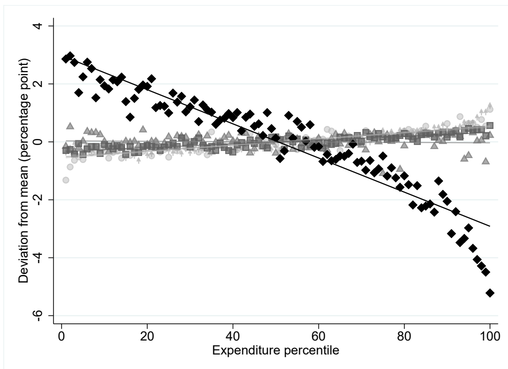
(c) CES

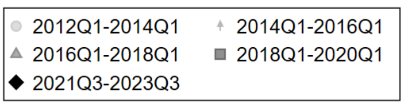


(d) Törnqvist



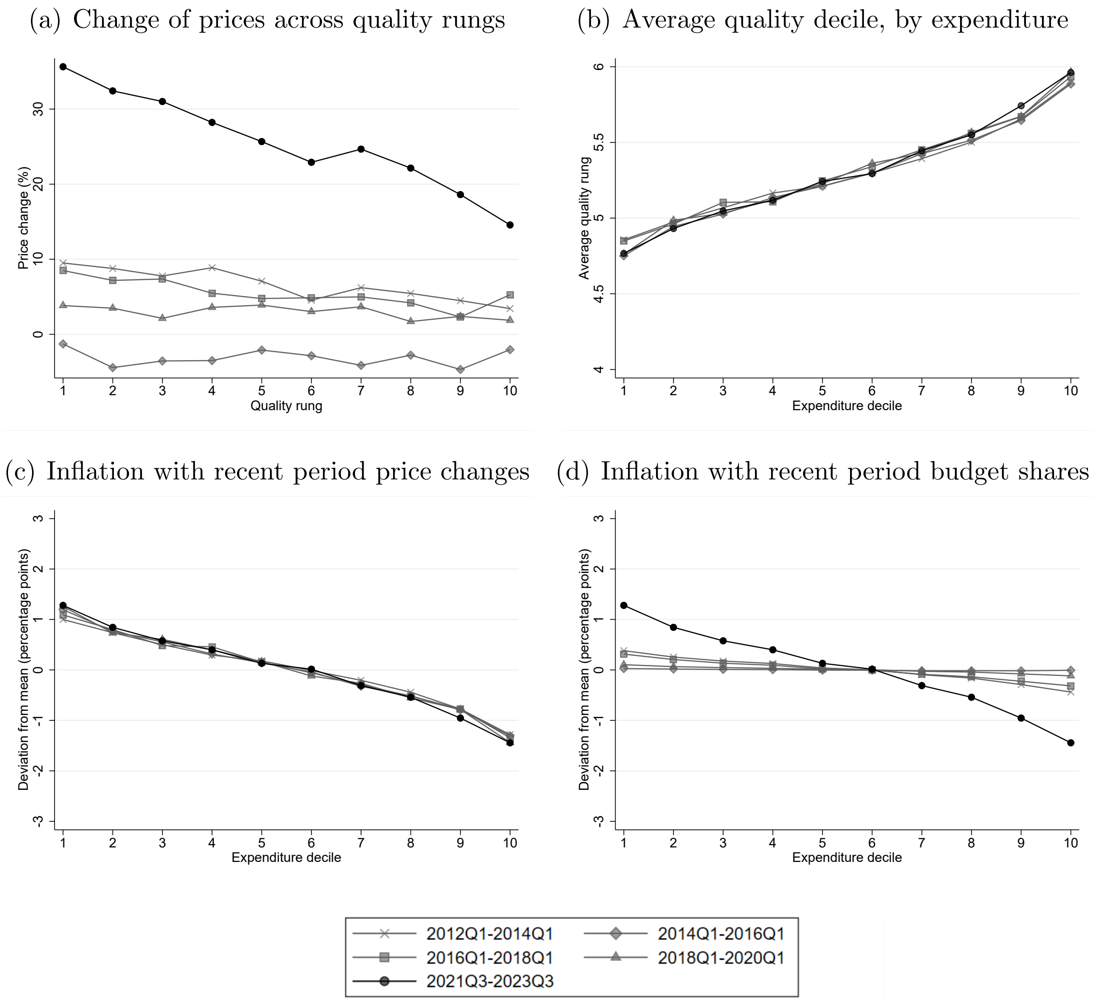
(e) Fisher





Notes: Authors' calculations using Kantar's Take Home panel (2012-2023). Figure plots the relationship between 9th quarter cumulative inflation and percentile of the expenditure distribution a household belongs to; there is a marker for each percentile and a line of best fit. We allocate households to expenditure percentiles based on their equivalized spending over the initial calendar year of the relevant nine quarter period. Cumulative inflation in panel (a) is based on a direct 2021Q3-2023Q3 comparison, and in panel (b) on a quarter-to-quarter chained comparison. All other panels are chained annually. The comparison figures in the paper are Figures 1(c) and 5(a) and 6(b).

Figure 13: *Inflation and quality rungs (adjusted for nonlinear pricing)*



Notes: Authors' calculations using Kantar's Take Home panel (2012-2023). In this figure quality rungs are defined by the procedure outlined in Appendix B. Panel (a) reports initial quarter-weighted increases in average price over the nine quarter period for products on each rung of the quality ladder. Panel (b) reports the average quality rungs of households' purchases by deciles of the expenditure distribution. Panels (c) and (d) report the relationship between cumulative inflation and expenditure decile in each nine quarter period, calculated using 2021Q3-2023Q3 price changes (panel (c)) and 2021Q3-2023Q3 spending shares (panel (d)). We allocate households to expenditure deciles based on their equivalized spending over the initial calendar year of the relevant nine quarter period. Cumulative inflation is computed using an annually chained Laspeyres index. The comparison figures in the paper are in Figure 4.