

# Did it ‘Really’ Happen? Cost of Living Inequality in Argentina, 2004–2018

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## Abstract

Measures of inequality typically rely on price indices formulated for a representative consumer, falsely assuming an identical consumption basket across households within the country. However, consumption patterns differ across households; hence, the changes in price levels might have different impacts on households at different points in the income distribution. To challenge this prevailing assumption and gain a more accurate comprehension of income inequality in Argentina from 2004 to 2018, I constructed income level-specific cost of living indices using the Argentinian Household Expenditure Survey. The results demonstrated that from 2004 to 2012, the poor experienced a higher increase in the cost of living compared to the richest group. Conversely, between 2012 and 2018, price changes displayed anti-rich behaviour. Considering the cost of living index differentials in the Gini coefficient calculations highlights that the price movements had an inegalitarian bias between 2004 and 2012, whereas the trend reversed in the 2012–2018 period.

**Keywords:** cost of living inequality, Gini coefficient, Argentina, QUAIDS.

**JEL Classification:** D13, D63, C81

## 1. Introduction

Inequality measures mostly rely on price indices designed for a representative consumer. This assumption might produce ideal indicators for inequality where an identical consumption basket is purchased by every household within the country. Nevertheless, Ernst Engel (1857) showed that this assumption is not held in the real world. He observed that households at different

points in the income distribution have different consumption patterns. Due to the differentials in consumption patterns, changes in relative prices or given price increases impact the households in the economy differently, and as such, households experience different cost of living indices (Muellbauer, 1974). Thus, neglecting the cost of living disparities between households can produce overestimated or underestimated income inequality. Therefore, constructing correct deflators to adjust the income of households is a crucial issue in measuring inequality (Crawford and Smith, 2002; Pendakur, 2002).

The very high levels of inflation experienced by Argentina make this problem particularly important in this context. The significance of this issue is heightened not only by the prevailing high inflation but also by the historical role that relative prices have played within the economy. The importance of relative prices in the Argentine economy is rooted in long-run distributional conflict over trade policies between the import-competing sector, including the working class, and the export sector since the beginning of the 20th century (Díaz Alejandro, 1970; O'Donnell, 1978; Gerchunoff and Alejandro, 1989; Debowicz and Segal, 2014). With his analysis, O'Donnell (1978) explained this conflict with the stop-go cycle model, which is a consequence of the wage-good effect<sup>1</sup>. He discussed that compared to other Latin American countries, Argentina had a peculiarity in the economy; Argentina's main export products were also wage-goods, which constituted the majority of workers' consumption basket. Therefore, the impact of a change in the relative price of wage-goods on labour consumption can be immediately observed. In the other Latin American countries, primary export products had less impact on the consumption basket of the working class; hence, the relative price of their consumption basket (O'Donnell, 1978).

This debate shows the critical role of relative prices of tradables to non-tradables in the historical distributional conflict in Argentina. Although this topic attracted attention in the literature, understanding the precise impact of this key role on income distribution requires a household-level analysis. Nevertheless, due to the data limitations, this paper is not able to analyse the direct impact of the relative prices of tradables versus non-tradables on income distribution in Argentina. Instead, this thesis aims to understand the inequality outcome of this historical wage-goods effect in Argentina by examining the relationship between prices and income inequality between 2004 and 2018.

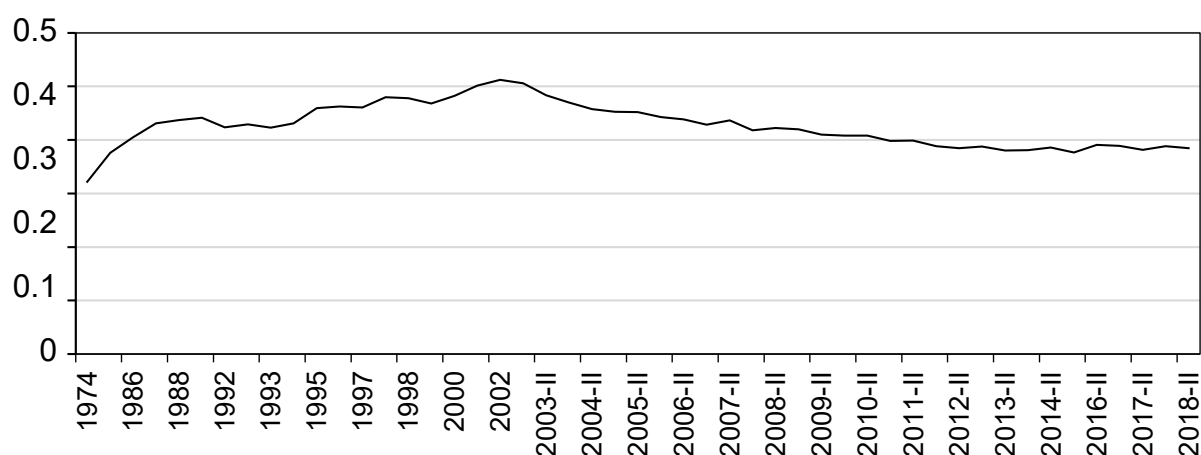
This paper focused on the period 2004, 2012 and 2018 for several reasons. First, as Figure 1 shows, income inequality in Argentina exhibited an increasing trend with upward and downward movements from 1974 to 2002. This trend reversed after 2002. Since the progress in income

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1 It should be noted that the “wage-good effect” term firstly used by Richardson (2009).

distribution coincided with commodity boom of the 2000s and the recovery era after the 2001 economic crisis in Argentina, many studies explained the decline in income inequality through these two phenomena (Cruces and Gasparini, 2008; Gasparini, Cruces and Tornarolli, 2011; Lustig, Lopez-Calva and Ortiz-Juarez, 2013; Judzik, Trujillo and Villafañe, 2017; Alvaredo, Cruces and Gasparini, 2018; Fernández and Messina, 2018; Fernández and Serrano, 2022). However, these studies have focused on patterns of nominal income inequality in Argentina, overlooking the distributional consequences of price changes and neglecting the historical wage-goods effect in Argentina.

**Figure 1: Gini Coefficient of Argentina, 2001-2018**



Source: Socio-Economic Database for Latin America and the Caribbean (*Centro de Estudios Distributivos, Laborales y Sociales, CEDLAS*) and The World Bank, August 2022 (latest version).

Secondly, during the Kirchner administration, the Argentine central bank heavily controlled the foreign exchange market to protect the peso from appreciating, and a stable and competitive exchange rate regime became the central macroeconomic policy of the era. Kirchner benefited from the undervalued exchange rate and the commodities boom to support domestic industry. Moreover, the administration followed *Peronista* tradition and supported the poor and working class through these policies, wage-goods subsidies and price agreements with businesses. In 2015, programmatic orientation significantly shifted from protectionism to trade liberalisation with the victory of right-wing *Propuesta Republicana*. Macri's economic stabilisation program was built on a trade-liberalisation policy and price increases in public services; therefore, the government revoked capital and exchange controls and raised public utility tariffs. (O'Farrell et al., 2022; Santarcangelo & Padín, 2021).

Finally, although energy subsidies were introduced to play a key role in protecting the most vulnerable population and urban working class from rising energy prices after the 2001 crisis, several studies have shown that these subsidies had negative distributional impacts due to significant inclusion errors. Their results show that these subsidies were distributed to all income groups and that the share of the richest households in total energy subsidies was approximately two times higher than the poorest households (Hancevic, Cont and Navajas, 2016; Lakner *et al.*, 2016; Giuliano *et al.*, 2020). Given these findings, one can expect that the poor might have experienced a higher increase in their cost of living than the richer households.

Given the reversed trend in income inequality, the exogenous shocks, policy changes, and, more importantly, the wage-goods effect, it is particularly critical to examine the distributional impact of price movements in Argentina between 2004 and 2018.

In this paper, I analyse the distributional impacts of price changes on income inequality in Argentina in 2004, 2012 and 2018 by constructing income-specific cost-of-living indices (COL) and assess how disparities in these indices across income groups translated into income inequality during the sample period. To do so, I employ the Laspeyres and Törnqvist indices as axiomatic approaches and the Quadratic Almost Ideal Demand System (QUAIDS) as a parametric approach. As a final step, to estimate the real Gini coefficient, I deflate the equivalised household incomes using the income-specific cost of living indices.

The analysis uses microdata from Argentinian Household Expenditure Surveys (*Encuesta Nacional de Gastos de los Hogares*) for 2004/05, 2012/13, and 2017/18, prepared by the National Institute of Statistics and Census of Argentina (INDEC). The Argentinian household expenditure survey does not record any data on prices; therefore, I construct unit values for each product by dividing the total expenditure on the product by the total quantity of purchased product. Unit values reflect the combination of quality and quantity decisions. A quality decision is income-dependent, and unit values do not hold quality constant (Gibson and Kim, 2013). Deaton (1988) discusses how richer households purchase higher-quality products, which increases the unit value higher for a given price level. In this scenario, one may conclude that poorer households face lower prices than other households. Hence, one may find that real income inequality is lower than nominal income inequality (Gibson and Kim, 2013). To deal with this quality bias, I calculate quality-adjusted unit values following Cox and Wohlgenant's (1986) approach. Then, I compute median unit values at the national level and assume that every household faces the same national median unit values within the country. Consequently, in my analysis, the differences in the cost of living across income groups stem from the variety in spending patterns of income groups.

The findings suggest considerable variation in the cost-of-living indices across income groups, especially between 2004 and 2012. The increase in the cost of living was biased against the poorest households and the lower-middle-income groups at the end of the commodity boom, being highest for the first decile. On the other hand, the top decile experienced the lowest increase in their cost of living. Calculations of subgroup CPIs and their contributions to the overall CPI show that the main reason behind this bias was the substantial increase in the price of food and beverages.

In contrast to the 2004–2012 period, the richest part of the population experienced the highest increase in the cost of living between 2012 and 2018, as the price indices of luxuries such as leisure and eating outgrew faster than average prices and necessity prices. In addition to these anti-rich price movements, dramatic rises in transport and communication and housing prices due to the rise in public utilities made a significant contribution to the cost of living index of the top decile.

Accounting for the cost of living differentials across income deciles reveals that real income inequality decreased at a slower rate than nominal income inequality from 2004 to 2012. The results for the latter period show that the real Gini coefficient in 2018 is lower than the nominal Gini coefficient, indicating that price movements had an equalising impact.

This paper contributes to two strands of the literature: first, this paper attempts to fill the gap in the literature on income inequality in Argentina by comparing nominal income inequality and real income inequality between 2004 and 2018. To the best of my knowledge, this study is the first attempt to calculate the cost of living inequality in Argentina and account for the cost of living differentials across income groups in Argentina's Gini Index estimations.

Secondly, this study extends the literature on inflation heterogeneity by focusing on cost-of-living differentials across income groups. Several papers in the literature examine the variation in inflation rates across income groups as well as across other sub-population groups such as age, household size, occupations, and urban and rural areas (Argente and Lee, 2021; Michael, 1979; Garner, Johnson and Kokoski, 1996; Amble and Stewart, 1994; McGranahan and Paulson, 2005; Hobijn and Lagakos, 2005; Taktek, 1998; Chiru, 2005a, 2005b; Murphy and Garvey, 2004; Crawford and Smith, 2002; Janský and Hait, 2016; Fessler and Fritzer, 2013; Gürer and Weichenrieder, 2020; Arndt, Jones and Salvucci, 2015; Goñi, López and Servén, 2006; AlAzzawi, 2016). In these works, the inflation rate across different groups was measured by matching the Consumer Expenditure Survey and the CPI price dataset, assuming that all households face the average price for each category of products.

Recent works challenged this standard assumption that every household faces the same prices by taking advantage of large scanner datasets collected at the consumer level (Kaplan

and Schulhofer-Wohl, 2017). Broda and Romalis (2008) find that most of the inflation heterogeneity comes from variation in prices paid for the same types of goods (almost 2/3 of the variation), not from variation in broadly defined consumption bundles (only 7 per cent of the variation). Jaravel (2019) investigated inflation inequality across income deciles in US using two datasets: the Consumer Expenditure Survey matched with CPI data and Nielsen scanner data, for the period 2004–2015. According to his findings, when matched data were used, the annual average inflation gap between the bottom quintile (2.20 per cent) and the top quintile (1.86 per cent) was 0.34 percentage points, whereas the gap between the bottom (1.87 per cent) and top quintiles (1.21 per cent) was 0.661 percentage points when based on Nielsen data. Through within-between decomposition, Jaravel demonstrated that most of this difference stems from within-category price differences, suggesting that inflation inequality is more pronounced in scanner data than in matched CPI data.

The rest of the paper is as follows: Section 2 provides a detailed discussion of methodology and data preparation. Section 3 introduces findings and discussion. Finally, section 4 concludes.

## 2. Methodology and Data Preparation

The cost-of-living index was introduced by Konüs (1939). He defined the cost of living as the monetary value of goods and services consumed in a certain period by the average household belonging to the given stratum of society. In the cost of living index theory, consumption of quantities represents the general state of the household's standard of living, and any standard of living may be attained by a different combination of quantities. He argues that if the general status of the standard of living of the household stays the same between the two periods of time, the "true cost of living index" can be obtained by dividing the cost of living at one period of time by the cost of living at the other period:

$$P_T = \frac{c(p_t, u_R)}{c(p_s, u_R)} \quad (1)$$

where  $c(p_t, u_R)$  is the cost function of the consumer in period  $t$  price vector  $p_t$  and the reference utility level  $u_R$ . Thus, the cost-of-living index (COLI) represents the relative change in the monetary cost of the consumption basket necessary for maintaining a certain standard of living.

In the literature, two main methods are used to construct the cost of living index: price indices and demand system estimations. The Laspeyres and Paasche indices are the most common price indices applied in the literature. However, these fixed consumption basket approaches fall short of capturing substitution bias: when consumers face price changes, they

are expected to alleviate the effects of price changes on their cost of living by adjusting their consumption patterns (Mackie and Schultze, 2002). To eliminate this bias in the COLI, I use the Törnqvist index as a main approach in this paper.

Diewert (1976) improved the Laspeyres and Paasche formulas without extending the information required:

$$\log P(p^1, p^0; T) = \sum_k 1/2 (w_k^1 + w_k^0) \log(p_k^1 / p_k^0) \quad (2)$$

where  $w^1$  and  $w^0$  are the budget shares in the current time and base time, respectively. Diewert (1976) illustrates that the Törnqvist index  $P(p^1, p^0; u^*)^2$  can be considered as a true index when the logarithm of the cost function is a quadratic form in the logarithms of prices and utility. Diewert (1976) also showed that the Törnqvist index is exact for a general non-homothetic translog cost function with the base of reference utility  $u^*$ . Hence, it allows for non-homothetic preferences, implying that the composition of the consumption basket is not independent of total expenditure or utility.

Although price indices provide several benefits, such as computational practicality and requiring readily observable price and demand vectors, the cost-of-living index can be accurately estimated if utility functions are known. Nevertheless, utility functions are not observable as prices and demand vectors. The standard solution to address this problem in the literature is to apply a demand system model, which allows to recover of cost functions from utility functions by producing required parameters (Deaton and Muellbauer, 1980; Balk, 1990; Schultze, 2003; Lluberas, 2018; Jaravel, 2021). To recover cost functions, I estimate the Quadratic Almost Ideal Demand System model, which was introduced by Banks, Blundell and Lewbel (1997). This model aims to construct a demand system permitting the more general Engel curve shapes. To do so, they define the indirect utility as:

$$\begin{aligned} V^h(p, m) &= \left[ \left( \frac{\ln m - \ln a(p)}{b(p)} \right)^{-1} + \lambda(p) \right]^{-1} \\ &= \left[ \frac{b(p)}{\ln m - \ln a(p)} + \lambda(p) \right]^{-1} \\ &= \frac{\ln m - \ln a(p)}{b(p) + \lambda(p) [\ln m - \ln a(p)]} \end{aligned} \quad (3)$$

2  $u^*$  is the reference utility level, which is the geometric mean of base utility and observed utility.

where  $\frac{\ln m - \ln a(p)}{b(p)}$  is the indirect utility function PIGLOG and  $\lambda(p)$  is a function of prices  $p$  differentiable and homogenous of degree zero; and  $\ln a(p)$  and  $b(p)$  are the price indexes from the AIDS model:

$$\log a(p) = \alpha_0 + \sum_k \alpha_k \log p_k + 1/2 \sum_k \sum_j \gamma_{kj}^* \log p_k \log p_j \quad (4)$$

$$\log b(p) = \log a(p) + \beta_0 \prod p_k^{\beta_k} \quad (5)$$

$$b(p) = \prod_i p_i^{\beta_i} = \exp(\sum_i \beta_i \ln p_i) \quad (6)$$

$$\lambda(p) = \sum_i \lambda_i \ln p_i \quad \text{where} \quad \sum_i \lambda_i = 0 \quad (7)$$

The above substituted into the QUAIDS indirect utility function provides:

$$V^h(p, m) = \left[ \left( \frac{\ln m - (\alpha_0 + \sum_k \alpha_k \ln p_k + 1/2 \sum_k \sum_j \gamma_{kj}^* \ln p_k \ln p_j)}{\prod_i p_i^{\beta_i}} \right)^{-1} + \sum_i \lambda_i \ln p_i \right]^{-1} \quad (8)$$

Hence, the cost function can be written as:

$$\log c(u, p) = \alpha_0 + \sum_k \alpha_k \log p_k + \frac{1}{2} \sum_k \sum_j \gamma_{kj}^* \log p_k \log p_j + \frac{u \prod_i p_i^{\beta_i}}{1 - u \sum_i \lambda_i \ln p_i} \quad (9)$$

$$\ln c(u, p) = \ln a(p) + \frac{ub(p)}{1 - u\lambda(p)} \quad (10)$$

When Shepard's lemma  $(\frac{\partial \log c(u, p)}{\partial \log p_i} = \frac{p_i q_i}{c(u, p)} = w_i)$  is applied to the above cost function, and  $u$  is substituted for the QUAIDS indirect utility function  $V(p, m)$ , the QUAIDS budget shares equations are obtained:

$$w_i = \alpha_i + \frac{1}{2} \sum_k \gamma_{kj}^* \ln p_k + \beta_i \left[ \ln \frac{m}{a(p)} \right] + \frac{\lambda_i}{b(p)} \left[ \ln \frac{m}{a(p)} \right]^2 \quad (11)$$

where expenditure  $x = m/a(p)$ .



## Data Preparation

I use the Argentinian National Household Expenditure Surveys (Encuesta Nacional de Gastos de los Hogares, ENGHo), which provide detailed data on households' expenditures and socio-economic variables. INDEC began conducting this cross-section survey in 1985 and has published five surveys (1985/86, 1996/97, 2004/05, 2012/13 and 2017/18) to date. These surveys are extensive; data were collected from 29,138 households in 2004, 20,944 households in 2012 and 21,547 households in 2018.

The main advantage of the Argentinean household expenditure survey (ENGHo) is that it includes disaggregated expenditure items (approximately 1000) at the household level and provides information on purchased quantities. The main drawback of the ENGHo is its limited rural coverage. The representativeness of the ENGHo varies across waves: the 2004–2005 wave was nationally representative, covering both urban and rural areas (INDEC, 2006), while the 2012–2013 and 2017–2018 waves restricted coverage to localities with 5,000+ and 2,000+ inhabitants, respectively (INDEC, 2013; INDEC, 2020), thus excluding some remote rural households—potentially among the poorest. According to World Bank estimates based on UN data, the rural population declined from 10.14% in 2004 to 8.13% in 2018, suggesting a limited impact on national estimates, though possible underrepresentation remains.

Also, to the best of my knowledge, no specific study exists on ENGHo's coverage of the richest, but evidence from the Household Income Survey (EPH) shows that surveys often underrepresent top incomes due to sampling limitations, low response rates, or exclusion of extreme values (Alvaredo, 2010; Alvaredo, Cruces & Gasparini, 2018; Alvaredo et al., 2022; World Bank, 2024). Thus, while these limitations do not invalidate the findings, they should be considered when interpreting inequality and cost-of-living results, as both rural undercoverage and top-income underrepresentation could slightly attenuate disparities.

As a first step, Following Deaton and Tarozzi (2005), I construct household-specific unit values due to a lack of actual price data:

$$v_{ih} = \frac{x_{ih}}{q_{ih}} \quad (12)$$

where  $v_{ih}$  is the household-specific unit value of the item  $i$ , and  $x_{ih}$  are the household's total expenditure on the purchased item and  $q_{ih}$  is the total quantity of the purchased item.

To detect the outliers in the unit values, in accordance with Cox and Wohlgemant (1986), observations with prices more than five standard deviations from the mean observed price are changed to missing values. Following Gibson and Rozelle (2005) and Cox and Wohlgemant

(1986), provincial and quarterly average unit values (at the elementary level) are calculated for each income decile to calculate sample means for missing unit values. The missing unit values are then imputed with these mean values. The remaining missing unit values (if there are any) are imputed with subregional and regional mean values, respectively.

The unit values represent households' purchase decisions; therefore, they display differences from one purchaser to another due to differentials in quality choice. To address this issue, I adjust the unit values for quality differences using Cox and Wolkstein's (1986) approach. I regress household characteristics on the log of unit values:

$$\log(v_{ih}) = \delta_i + c_h' k_i + \epsilon_{ih} \quad (13)$$

where  $\delta_i$  is constant,  $\epsilon_{ih}$  is the error term, and the vector  $c_h$  includes the education level, age and gender of the head of the household, household income, household size, and seasonal and regional dummies. Then, the vector of quality-adjusted unit values is calculated:

$$ph = \exp(\hat{\delta} + \hat{\epsilon}_h) \quad (14)$$

The outliers in total expenditures are detected using graphical methods following Deaton and Zaidi (2002). Ten different sub-datasets are created for each income decile in each period, and then one-way plots of total expenditures are drawn to detect outliers in each decile. Together with these outliers, the households with zero total expenditures are dropped.

## Aggregation of Budget Shares

As the number of parameters required to run a full system of demand rises with the square of the number of expenditure items, a certain level of aggregation becomes essential to applying a demand system model. I aggregate more than a thousand expenditure items into eleven consumption bundles according to the criteria of the Classification of Individual Consumption by Purpose (COICOP, 1999): food and beverages<sup>3</sup>, alcohol and tobacco, clothing and footwear,

3 In this paper, the consumption of household-produced goods, which refers to all primary products resulting from agricultural, livestock, farming, hunting, forestry, fishing, or mining activities, was included in the cost-of-living index calculations. According to the Household Expenditure Survey Handbook (ENGHO 2017–2018, Informe de Gastos), for the purposes of valuation, the retail price at which comparable products are sold is recorded as an expense, with households requested to provide an estimate of the equivalent cash price. In such cases, the value of any inputs acquired by the household for the production of these goods is not included in the recorded amount (INDEC, 2020). Descriptive statistics show that household-produced goods account for only 0.10% of total expenditure in 2017/2018 and 0.57% in 2004/2005.

house maintenance, housing, leisure, health, miscellaneous goods and services, finance and insurance, eating out and transport and communication. I exclude education, tourism and private transport consumption groups as INDEC did not collect any information on quantities in the 2004/2005 and 2012/2013 surveys for these commodity groups. These goods are generally regarded as luxury or high-income elastic items, and an examination of budget share patterns confirms that these categories are systematically more important for high-income households. Hence, their omission could result in a mild downward bias in the measured inequality, as the index would be missing categories whose relative price changes could disproportionately affect high-income households' cost of living. However, their aggregate budget shares remain small in comparison with essential goods such as food (43.6% in the poorest vs. 20.9% in the richest) or housing (12.2% vs. 14.6 %). Given these low shares, the magnitude of the bias is expected to be limited, although the direction of the effect—towards understatement of inequality—is clear.

Additionally, since my analysis covers fifteen years between 2004 and 2018, different goods and services are included/excluded in the household expenditure survey over time. Therefore, following Almås and Kjelsrud (2017), the items that do not appear in every survey are dropped from the dataset for the sake of consistency.

It is crucial to note that INDEC excludes imputed rents from the Argentine CPI basket. Although the ENGHo enumerators ask the households to estimate the rental values of owner-occupied dwellings, the estimated rental values are excluded from the cost of living calculations in this paper due to the reliability problem.

Another challenge I faced was misreported quantity information on utility (water, gas, and electricity) expenditures. Due to a lack of information, I exclude water and gas expenses from my calculations. To solve this misreporting problem in electricity expenses, I replace the unit values with electricity prices collected from three leading Argentinian distributors *Empresa Distribuidora y Comercializadora Norte* (EDENOR), *Empresa Distribuidora de Electricidad del Sur* (EDESUR), and *Empresa Distribuidora La Plata* (EDELAP).

After these consistency checks, I calculate eleven aggregated budget shares to obtain the QUAIDS budget shares equations given in 4:

$$w_j^h = \frac{\sum p_i^h q_i^h}{\sum p^h q^h} \quad (15)$$

where  $w_j^h$  is the budget share of the  $j$ th ( $i = 1, \dots, 11$ ) aggregated consumption bundle in the total consumption basket for each household,  $p_i^h, q_i^h$  is the total expenditure of  $i^{\text{th}}$  component of the aggregated consumption bundle and  $\sum p^h q^h$  is the total expenditure of the household.

## Aggregation of Prices for the QUAIDS Model

After computing the budget shares of the eleven aggregated consumption bundles, single prices for each aggregated bundle are constructed using the weighted geometric means for each household:

$$\ln P_{j,h} = \sum_i^n w_{i,j,h} \ln p_{i,j} \quad (16)$$

where  $\ln P_{j,h}$  is a group-level aggregated price for each household,  $w_{i,j,h}$  is the budget share of the expenditure item in its aggregated consumption bundle for each household (for instance, the weight of pink lady apples in the food and beverage bundle), and  $p_{i,j}$  is the median unit value of the expenditure item (Abdulai, 2002; Browning, Chiappori and Lewbel, 2013).

Calculating single prices for aggregated commodity bundles is not a problem-free approach due to the large amount of zero consumption in the dataset. If the  $h^{th}$  household did not consume any components of the aggregated bundle during the survey week, the price of the aggregated commodity bundle would be zero for the household. As I choose the median household to construct the income-specific cost of living index, having zero values for aggregated prices makes it impossible to track the changes in the income-specific cost of living index. Therefore, I replace these zero values (if there are any) with the income-specific quarterly and provincial mean aggregated prices to deal with this issue.

One of the biggest challenges in the demand system analysis was the large amount of zero consumption. To deal with the censored dependent problem, I apply Heien and Wesseils's (1990) two-step approach. I provide details on this approach in Appendix B.

## Data Preparation for The Törnqvist Price Index

The price indices at each higher level of aggregation can be estimated on the basis of weights and prices for its components, that is, the lower-level price indices (Graf, 2020). The higher-level indices are calculated simply as weighted averages of the lower-level indices. Following this approach, I first create price indices for each sub-aggregated consumption bundle that is constructed following the criteria of COICOP. Then, these sub-group price indices are aggregated together using their budget shares in the total consumption basket up to the Overall Consumer Price Index.

The weights that are used in the calculation of sub-aggregated price indices are the average weights of the components of the sub-groups of the CPI. For instance, white rice belongs

to the food and beverage sub-group, and the weight of white rice in a household's food price index is calculated by dividing the total expenditure of the household on white rice by the total expenditure of the household on food. The average weight of white rice for the poorest decile was computed by averaging the poorest households' budget share of white rice within the food and beverage consumption bundle.

The sum of the average weights of the components should be equal to 1. However, in the analysis, the use of the average weights of the components in sub-group price indices violates the adding-up rule in demand theory due to the large amount of zero expenditure. To deal with the adding-up violation, I calculate plutocratic weights for each income group by creating super households, which are achieved by adding up the expenditures on the component of all households who belong to the same income decile. The result is then divided by the households' total expenditure on sub-groups:

$$w_{i,j}^d = \frac{\sum p_{i,j}^{h,d} q_{i,j}^{h,d}}{\sum p_j^{h,d} q_j^{h,d}} \quad (17)$$

where  $w_{i,j}^d$  is the average budget share of the  $i^{\text{th}}$  ( $i = 1, \dots, n$ ) component in  $j^{\text{th}}$  ( $j = 1, 2, \dots, 11$ ) sub-group consumption basket for  $d^{\text{th}}$  income decile,  $\sum p_{i,j}^{h,d} q_{i,j}^{h,d}$  is the total expenditure of  $d^{\text{th}}$  decile on components, and  $\sum p_j^{h,d} q_j^{h,d}$  is the total expenditure of  $d^{\text{th}}$  decile on sub-group consumption basket.

Then, the sub-group price indices are calculated with the following formula:

$$\log P_j^d(p_j^1, p_j^0; T) = \sum 1/2(w_i^{1,d} + w_i^{0,d}) \log(p_i^{1,d} / p_i^{0,d}) \quad (18)$$

where  $p_j^d$  is the  $j^{\text{th}}$  ( $j = 1, 2, \dots, 11$ ) sub-group price index for  $d^{\text{th}}$  decile,  $w_i^{1,d}$ , and  $w_i^{0,d}$  are the income-specific average weights of the component within  $j^{\text{th}}$  sub-group consumption bundle in observed year and base year, respectively.  $p_j^{1,d}$  and  $p_j^{0,d}$  are the median unit values of the component in the observed year and base year, respectively.

After calculating the sub-group price indices, I aggregate the sub-group price indices using their expenditure shares in the total consumption basket up to the income-specific Overall Consumer Price Index.

Members of the households have different needs based on their gender, age, and other demographic characteristics. Considering these facts, I adjust households' disposable incomes by applying equivalence scales considering the Argentinian households' size and composition. Following Deaton and Zaidi's (2002) advice, I apply the arbitrary approach to equalise household incomes. The formula is given below (CEDLAS, 2014):

$$E = (A + \alpha_1 K_1 + \alpha_2 K_2)^\theta \quad (19)$$

where  $A$  is the number of adults (15 – 99+),  $K_1$  is the number of children under 5 years old, and  $K_2$  is the number of children between 6 and 14. Parameters allow for different weights for children and adults, and  $\theta$  regulates the degree of a household's economies of scale. Following Deaton and Zaidi (2002) and CEDLAS (2014), I set  $\alpha_1 = 0.5$  and  $\alpha_2 = 0.75$ , and  $\theta = 0.9$  for Argentinian households.

### 3. Findings and Discussion

As the sample period covers both the commodity boom era and the post-commodity boom era, the distributional impact of relative price changes may show noticeable differences in these periods. Therefore, I examine the changes in the COLIs in the two eras separately. To calculate the COLI for the period 2004–2012 and 2004–2018, I used 2004 as the base year. For the post-commodity boom era (2012–2018), I used 2012 as the base year. The results are given in Table 1.

The Törnqvist index shows that the COLI monotonically decreased as income levels increased between 2004 and 2012, revealing that the price movements were anti-poor at the end of the commodity boom in Argentina. The top and middle-income groups experienced lower increase rates in the cost of their consumption baskets than the average and the rest of the Argentinian households.

**Table 1: The Cost of Living Indices for Equivalized Income Deciles, 2004–2018.**

| Deciles                   | Törnqvist |           |           | Laspeyres |           |           |
|---------------------------|-----------|-----------|-----------|-----------|-----------|-----------|
|                           | 2004—2012 | 2012—2018 | 2004—2018 | 2004—2012 | 2012—2018 | 2004—2018 |
| <b>1st</b>                | 4.05      | 4.51      | 19.06     | 4.54      | 6.11      | 26.34     |
| <b>2nd</b>                | 3.97      | 4.67      | 20.07     | 4.42      | 6.61      | 26.94     |
| <b>3rd</b>                | 3.95      | 4.72      | 19.96     | 4.46      | 6.90      | 26.42     |
| <b>4th</b>                | 3.90      | 4.74      | 19.76     | 4.51      | 6.58      | 25.64     |
| <b>5th</b>                | 3.86      | 4.76      | 19.73     | 4.40      | 6.73      | 25.14     |
| <b>6th</b>                | 3.85      | 4.78      | 19.65     | 4.44      | 6.71      | 25.45     |
| <b>7th</b>                | 3.83      | 4.81      | 19.58     | 4.42      | 6.77      | 24.59     |
| <b>8th</b>                | 3.79      | 4.81      | 19.70     | 4.46      | 6.89      | 25.10     |
| <b>9th</b>                | 3.74      | 4.91      | 19.75     | 4.44      | 6.87      | 25.46     |
| <b>10th</b>               | 3.71      | 5.01      | 20.09     | 4.54      | 7.50      | 26.33     |
| <b>National (Average)</b> | 3.84      | 4.82      | 19.83     | 4.47      | 6.86      | 25.68     |

Source: Author's own calculations. The dataset is obtained from ENGHo 2004/2005, 2012/2013 and 2017/2018.

Unlike the 2004-2012 period, prices behaved in an anti-rich manner as the richest part of the population experienced the highest increase in the cost of living between 2012 and 2018. Over the sample period, the rate of increase in the COLI for the top and second income deciles declined noticeably faster than for the rest of the population. By implication, the richest households in Argentina are expected to have faced the most significant loss in purchasing power of money due to the sharp rise in prices between 2004 and 2018.

The second panel provides the results of the Laspeyres-based COLI. Consistent with the price index theory, the Laspeyres index outpaced the Törnqvist-based COLI for each income decile; hence, it overstated the cost of living index of income groups in each sample period. Despite this overstatement, the Laspeyres index tells the same story as the Törnqvist index for both the 2012-2018 period and the full sample period, indicating that the richest households experienced the highest cost of living index.

The estimation of Laspeyres indices is undertaken to make comparison possible with the official CPI figures. However, at the beginning of 2007, official CPI<sup>4</sup> calculated by INDEC underwent a series of methodological alterations which significantly undermined its credibility (Cavallo and Bertolotto, 2016). Additionally, from December 2015 to April 2016, INDEC stopped publishing official CPI due to a statistical emergency. Therefore, this paper benefits from an alternative CPI<sup>5</sup> produced by Cavallo and Bertolotto (2016), which offers a credible measure of CPI.

Comparing Table 1 and Figure 2 indicates that the calculated CPI in this paper showed a similar increase to the alternative CPI between 2004 and 2012, whereas the calculated CPI displayed a higher increase than the alternative CPI in 2018. This divergence could be partially explained by the products involved in the alternative CPI. Cavallo and Bertolotto's CPIs are based on online prices of two Argentine retailers, which contain only food, beverages, and household products, while the CPI computations in this paper cover a broader range of goods and services.

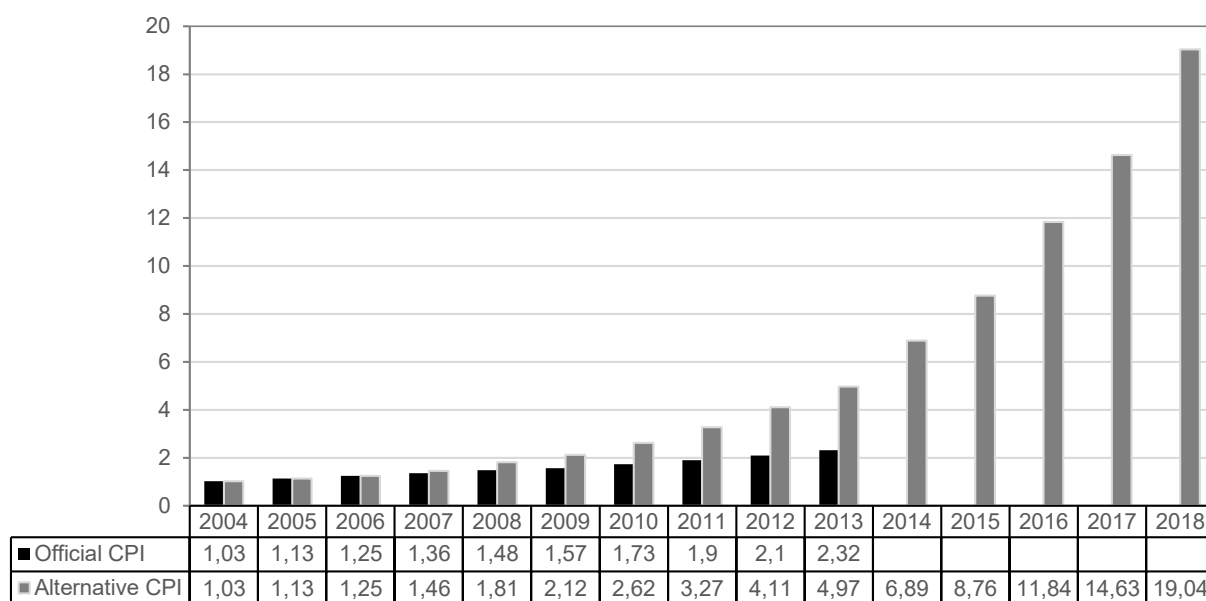
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4 The data from 2014 to 2018 are not available due to the reliability of the Official CPI. From 1943 to 2013, the CPI was only collected in Greater Buenos Aires. From 2014 to October 2015, not only was the base changed (from 2008=100 to 2014 = 100), but the National Urban CPI (IPCNu) was also collected. From November 2015 to March 2016, due to the statistical emergency, INDEC stopped publishing CPI. From April 2016 to December 2016, the CPI was resumed for Greater Buenos Aires. From December 2016 to the present, the National CPI has been compiled, publishing the national total and by statistical region. Thus, the information from the different databases cannot be combined. Not only were there database changes, but the universe of the survey also changed.

5 Cavallo and Bertolotto (2016) calculated CPIs from 1943 to the present by replacing the discredited official inflation rates (from 2007 to 2016) with an index built by PriceStats. The PriceStats is a private company that collects and uses online price data to estimate inflation indicators in many countries. A weakness of the collected online prices is that the data does not cover all the products involved in the official CPI (Cavallo and Bertolotto, 2016).

As will be widely discussed in the next section, the increase in CPI between 2004 and 2012 was mainly driven by food inflation, while cuts in public utilities, along with food and beverages, contributed significantly to the CPIs between 2012 and 2018. Hence, the lower increase in the alternative CPI could partly be explained by the limited coverage of the alternative CPI.

**Figure 2: The Official CPI and The Alternative CPI, 2014—2018.**



Source: INDEC (2014) and Cavallo and Bertolotto (2016).

### 3.1 Sub-Group Consumer Price Indexes and Inflation Decomposition

To understand the manner of price movements, I took one step further and decomposed the total inflation rate by using lower-level expenditure items. To analyse how the sub-group CPIs contribute to income-level-specific inflation rates, I applied the OECD decomposition formula where budget shares are kept fixed for more than a year:

$$Contrib_j^y = W_j^{y_0} \times \frac{p_j^y - p_j^{y-1}}{p_{TOT}^{y-1}} \quad (20)$$

where  $W_j^{y_0}$  is the budget share of COICOP component  $j$  in year  $y_0$  (base period),  $p_{TOT}^{y-1}$  is the overall price index in year  $y-1$  and  $p_j^y$  is the price index of COICOP component  $j$  in year  $y$ . In this formula, the contribution of a commodity group to the overall inflation rate is independent of price changes of other commodity groups, and the sum of the contributions of all commodity groups should be equal to the overall inflation rate (OECD, 2022).



Since the contribution formula requires base year weights, I used the Laspeyres index to compute the sub-group CPIs and their contribution to the inflation rate in this part. Table 2 reports the sub-group CPIs and their contribution to the overall CPI for the first decile and tenth decile.

Table 2 illustrates that the most significant contributor to the inflation rate of the poor was the food and beverages bundle in each period. This trend can be attributed to both the high growth rate of food and beverage CPI and the substantial share of these expenditures in the consumption basket of the poor. The poorest 10 per cent of Argentinean households spend more than half of their budget on food and beverages<sup>6</sup>; therefore, even a slight change in food prices can notably impact their cost of living.

**Table 2: The Subgroup CPIs and Contributions – First and Tenth Income Decile**

|                                         | First Income Decile |           |           |               |           |           | Tenth Income Decile |           |           |               |           |           |
|-----------------------------------------|---------------------|-----------|-----------|---------------|-----------|-----------|---------------------|-----------|-----------|---------------|-----------|-----------|
|                                         | Subgroup CPI        |           |           | Contributions |           |           | Subgroup CPI        |           |           | Contributions |           |           |
|                                         | 2004–2012           | 2012–2018 | 2004–2018 | 2004–2012     | 2012–2018 | 2004–2018 | 2004–2012           | 2012–2018 | 2004–2018 | 2004–2012     | 2012–2018 | 2004–2018 |
| <b>Food and Beverages</b>               | 5.19                | 4.52      | 22.08     | 2.17          | 1.71      | 10.89     | 5.15                | 4.97      | 23.47     | 0.86          | 0.98      | 4.64      |
| <b>Tobacco and Alcohol</b>              | 3.69                | 20.35     | 86.69     | 0.08          | 0.40      | 2.63      | 3.98                | 7.46      | 26.21     | 0.07          | 0.13      | 0.61      |
| <b>Clothing and Footwear</b>            | 3.29                | 6.12      | 26.94     | 0.22          | 0.55      | 2.47      | 2.88                | 7.43      | 27.10     | 0.20          | 0.70      | 2.73      |
| <b>Home Maintenance</b>                 | 4.17                | 6.84      | 22.19     | 0.18          | 0.38      | 1.22      | 4.45                | 10.06     | 21.02     | 0.22          | 0.59      | 1.29      |
| <b>Housing</b>                          | 3.36                | 13.09     | 44.37     | 0.22          | 0.81      | 4.03      | 3.70                | 11.40     | 32.87     | 0.29          | 1.39      | 3.45      |
| <b>Leisure</b>                          | 4.41                | 7.56      | 24.15     | 0.11          | 0.34      | 0.78      | 4.30                | 8.52      | 22.90     | 0.32          | 0.68      | 2.12      |
| <b>Health</b>                           | 5.01                | 5.77      | 20.26     | 0.16          | 0.15      | 0.77      | 5.87                | 7.18      | 35.30     | 0.29          | 0.29      | 2.03      |
| <b>Miscellaneous Goods and Services</b> | 4.68                | 5.43      | 22.02     | 0.17          | 0.20      | 0.99      | 5.10                | 7.55      | 27.69     | 0.22          | 0.32      | 1.46      |
| <b>Finance and Insurance</b>            | 5.89                | 3.92      | 21.18     | 0.01          | 0.01      | 0.06      | 7.22                | 3.70      | 19.83     | 0.25          | 0.06      | 0.76      |
| <b>Eating out</b>                       | 6.24                | 8.27      | 37.60     | 0.08          | 0.09      | 0.54      | 5.88                | 8.75      | 41.05     | 0.45          | 0.43      | 3.71      |
| <b>Transport and Communication</b>      | 2.97                | 5.27      | 14.99     | 0.14          | 0.47      | 0.96      | 3.44                | 6.79      | 18.07     | 0.36          | 0.94      | 2.52      |
| <b>Overall CPI</b>                      | 4.54                | 6.11      | 26.34     | 3.54          | 5.11      | 25.34     | 4.54                | 7.50      | 26.33     | 3.54          | 6.50      | 25.33     |

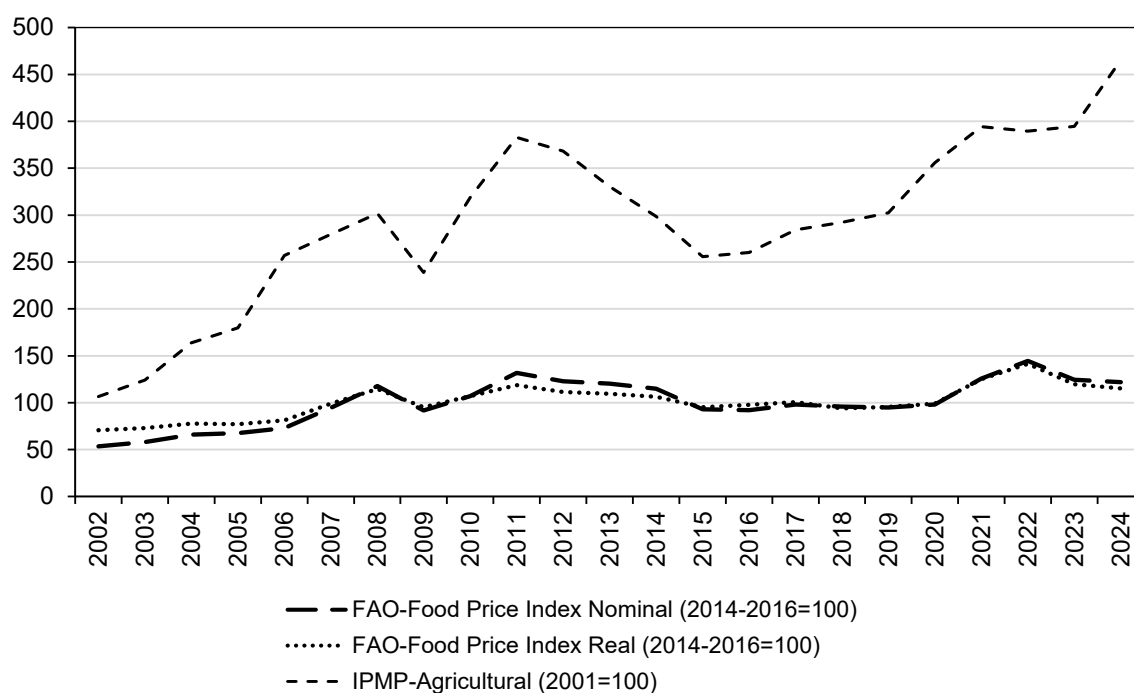
Source: Author's own calculations. The dataset is obtained from ENGHo 2004/2005, 2012/2013 and 2017/2018.

6 See Appendix A, Table A.1.

Between 2004 and 2012, the food and beverages CPI grew faster than the other necessity groups and the average inflation rate. This steep rise could be driven by the rise in aggregated demand after the 2001 crisis and international food prices during the 2000s commodity boom. As illustrated in Figure 3, international food prices started increasing after December 2004, peaking in the last quarter of 2012. The Kirchner government responded to rising food prices through export taxes, price controls and price subsidy mechanisms (Gallacher and Lema, 2014; Damill, Frenkel and Rapetti, 2015). However, the results show that food inflation remained higher than the average inflation, which means that the government's policies fell short of protecting the urban poor from rising food prices.

Although there was no return to early 2000s price levels after 2012, food prices exhibited a significant decline until 2021. This decline may have contributed to the decreasing importance of the food and beverages group in the overall CPI between 2012 and 2018.

**Figure 3: International Food Price Index, 2002–2024**



Source: The IPMP dataset<sup>7</sup> is taken from the Central Bank of Argentina (*Banco Central de la República Argentina*). The FAO Food Price Index is taken from the Food and Agriculture Organisation of the United Nations<sup>8</sup>.

7 The IPMP (The Commodity Price Index) tracks the prices of most representative basic agricultural commodities (barley, beef, corn, wheat, soybean, soybean pellets and soybean oil), weighted by their share on total exports. The index is calculated using the available prices in the commodity markets of the US, Europe and Asia.

8 The FAO Food Price Index consists of the average of five commodity group indices: meat, dairy, cereals, vegetables and oils, and sugar, which are weighted by the average export shares of each group for 2014–2016.

From 2012 to 2018, the housing CPI grew faster than the other groups and the overall CPI, due to substantial increases in utility prices. While electricity prices did not show a considerable change between 2004 and 2012, they sharply increased after 2012 due to the rise in public utility tariffs. After the 2001 economic crisis in Argentina, the Duhalde government froze the public utility tariffs to protect the purchasing power of Argentineans in an inflationary environment (González, 2009; Haselip and Potter, 2010; Bril-Mascarenhas and Post, 2015; Lakner et al., 2016). After the poor economic situation was over, the cost of the subsidies rapidly increased due to the surge in international energy prices and the rising demand for subsidised goods. Consequently, energy subsidies became a primary factor behind the growth of public expenditure; its share in GDP increased from 0.3 per cent to 3.5 per cent between 2005 and 2014 (Giuliano et al., 2020). In December 2015, Macri won the general election, and a new administration declared a series of measures to raise public utility tariffs, leading to a spike in electricity prices. Since the bottom decile devoted a significant amount of their housing budget to electricity bills<sup>9</sup>, this steep increase observed after 2012 may have contributed to the soaring housing inflation rate between 2012 and 2018.

Additionally, as with other public utilities, prices of bottled gas (LPG cylinders) significantly surged between 2012 and 2018. This increase particularly harmed poor households, who devoted an average of 21 per cent of their housing budget to bottled gas, as many poor households live in areas that are not connected to gas networks (Lakner et al., 2016). Subsequently, this amount is only 1 per cent for the richest decile.

As in the poorest decile, the food inflation rate was a key driver of the cost-of-living increases for the wealthiest households in 2012. Although the richest households allocated an average of 24 per cent of their income to food, less than half of the poorest food share, food and beverage was still the largest expenditure share for the top decile. As a result, this commodity group was the most significant contributor to the rich's inflation rate in 2012.

Between 2012 and 2018, the major driver of inflation for the richest population was the severe increase in housing CPI, mainly due to cuts in energy subsidies. Several studies in the literature show that the share of the richest households in total energy subsidies was approximately two times higher than the poorest households due to inclusion errors (Hancevic, Cont and Navajas, 2016; Lakner et al., 2016; Giuliano et al., 2020). Hence, cuts to energy subsidies were reflected in the housing price index of the top decile, and housing CPI displayed the highest

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9 An average of 8.48 per cent of the poorest households paid rent during the sample period, while for the richest decile, this average was approximately three times higher. Due to the low tenancy rate, the budget share of rent is lower than the budget share of utilities in the housing consumption basket of both income groups.

increase rate from 2012 to 2018. Since the wealthiest households allocated an average of 47 per cent of their housing budget shared to electricity bills, housing CPI emerged as the primary contributor to the wealthiest households' inflation rate.

Table 2 shows that the subgroup CPI for luxury goods and services, such as eating out and leisure, grew faster than the other necessities (except housing CPI) and overall CPI, particularly for the top decile, between 2012 and 2018. Given that richer households spend more on these expenditure groups, the higher cost of living index observed for the richest households can also be attributed to the rise in the relative price of luxury goods and services.

### 3.2 QUAIDS-Based Cost of Living Index

As discussed in the methodology section, the cost of living index can be precisely calculated if utility functions are known. By applying the QUAIDS method, I obtain utility functions and recover the expenditure functions for each year by using the median prices, the median expenditure level and the base utility level that was set as the utility level of the representative households (median consumers). After calculating the expenditure functions for each decile, the cost-of-living index was obtained by dividing the cost of living for the observed year by the cost of living for the base year.

Table 3 reports the total expenditures of median households and the true cost of living indices for each income decile during the sample period. The total expenditure shows the amount needed to attain the base utility level.

Even though the magnitude of the increase rate of COLI shows differences compared to Törnqvist-COLI, both measures tell the same story about the changes in the cost of living across income groups. From 2004 to 2012, the increase in the COLI was biased against the lower income groups at the end of the commodity boom, being highest for the poorest decile. On the other hand, the top decile experienced the lowest increase in their cost of living. The reverse situation was observed between 2012 and 2018; in this period, the true cost-of-living index was strongly biased against the richest Argentinian households, whereas the lower-middle-income and middle-income groups experienced the smallest increase in their cost of living in 2018. For the full sample period, the QUAIDS-COLI shows that the richest households experienced the highest increase in their cost of living.

**Table 3: The Cost of Living Indices Estimated by the QUAIDS Model**

| Income Deciles   | Total Expenditure |          |           |   | True Cost of Living Index |           |               |
|------------------|-------------------|----------|-----------|---|---------------------------|-----------|---------------|
|                  | 2004              | 2012     | 2018      |   | 2004–2012                 | 2012–2018 | Sample Period |
| 1 <sup>st</sup>  | 367.75            | 1,823.26 | 8,319.53  | 1 | 4.96                      | 3.85      | 22.62         |
| 2 <sup>nd</sup>  | 494.37            | 2,439.87 | 11,264.70 | 1 | 4.94                      | 3.54      | 22.79         |
| 3 <sup>rd</sup>  | 592.32            | 2,931.10 | 13,147.34 | 1 | 4.95                      | 3.50      | 22.20         |
| 4 <sup>th</sup>  | 660.58            | 3,111.61 | 15,239.84 | 1 | 4.71                      | 3.75      | 23.07         |
| 5 <sup>th</sup>  | 702.90            | 3,259.05 | 15,879.80 | 1 | 4.64                      | 3.66      | 22.59         |
| 6 <sup>th</sup>  | 797.10            | 3,731.04 | 19,313.78 | 1 | 4.68                      | 3.86      | 24.23         |
| 7 <sup>th</sup>  | 915.21            | 4,180.82 | 21,717.95 | 1 | 4.57                      | 3.87      | 23.73         |
| 8 <sup>th</sup>  | 1,077.54          | 5,157.21 | 27,133.79 | 1 | 4.79                      | 3.93      | 25.18         |
| 9 <sup>th</sup>  | 1,294.51          | 5,929.27 | 31,506.77 | 1 | 4.58                      | 4.17      | 24.34         |
| 10 <sup>th</sup> | 1,804.82          | 7,091.65 | 47,572.54 | 1 | 3.93                      | 5.76      | 26.36         |

Source: Author's own calculations. The dataset is obtained from ENGHo 2004/2005, 2012/2013 and 2017/2018.

The differentials in the increase rates of QUAIDS-COLI and Törnqvist COLI can be explained by methodological differences in creating aggregated prices and the weights used. Firstly, as extensively discussed in the data preparation section, I used plutocratic weights by creating super-households to construct Törnqvist and Laspeyres indices, while I used democratic weights to compute aggregated prices for the QUAIDS model. Although the indices and aggregated prices are income-specific, plutocratic weights tend to reflect consumption patterns of higher-income households. In contrast, democratic weights display the consumption pattern of the average household. Calculating the weights in different ways changes the importance of items in the consumption basket due to the large amount of zero consumption. This impact directly changes the value of aggregated prices. It is crucial to note that the use of plutocratic weights in the QUAIDS model caused a multicollinearity problem between prices; hence, the QUAIDS model produces no parameters.

Secondly, axiomatic and parametric approaches require different methods to deal with zero consumption. Addressing this problem in the demand system models necessitates advanced econometric techniques, such as two-step estimators or Tobit models. In the context of the axiomatic approach, imputation methods can solve the issue of zero consumption, especially when combined

with plutocratic weighting. Thirdly, the traditional demand theory presumes homogeneous products with a single price. This assumption may introduce a production bias in results since the reality is more complex (Davis, 1997). Although the production bias is inherent in all approaches to some extent, the impact may be more severe for the QUAIDS model due to the high aggregation level.

### 3.3 Real Income Inequality

This section aims to use the income-specific COLIs to analyse how the cost-of-living differentials across income groups contributed to the income distribution in Argentina during and after the 2000s commodity boom. To examine the effect, I compute a 'real Gini coefficient' by deflating households' equivalised income by income level-specific cost of living index:

$$Real\ Income_{d^{0,t}} = \frac{Nominal\ Income_{d^{0,t}}}{Cost\ of\ Living\ Index_{d^{0,t}}} \quad (21)$$

where  $d$  represents the income decile, 0 is the reference period, and  $t$  is the current period. For the nominal income variable, I use the net total income of households, including labour income and non-labour income. Labour income consists of income from the main occupation, secondary occupation and previous occupations. Non-labour income is the sum of pensions, capital income, money transfer and valorisation of the production made for home consumption.

Table 4 reports on the evolution of income inequality and real income inequality over fourteen years. In the first panel, I deflate equivalized incomes by QUAIDS-COLI and in the second panel, I use the Törnqvist index to adjust equivalized household incomes. As is consistent with the recent literature on income inequality in Argentina (Gasparini and Cruces, 2010; Gasparini, Cruces and Tornarolli, 2011; Lustig, Lopez-Calva and Ortiz-Juarez, 2013; Judzik, Trujillo and Villafañe, 2017; Alvaredo, Cruces and Gasparini, 2018; Fernández and Messina, 2018; Fernández and Serrano, 2022), I find that nominal income inequality decreased during the 2000s commodity boom, but started increasing again after the boom ended.

**Table 4: Real Income Inequality in Argentina, 2004–2018**

| Years       | Adjusted with True Cost of Living Index |                       | Adjusted with Törnqvist Index |                       |
|-------------|-----------------------------------------|-----------------------|-------------------------------|-----------------------|
|             | Nominal Gini Coefficient                | Real Gini Coefficient | Nominal Gini Coefficient      | Real Gini Coefficient |
| <b>2004</b> | 0.467                                   | 0.467                 | 0.467                         | 0.467                 |
| <b>2012</b> | 0.398                                   | 0.432                 | 0.398                         | 0.408                 |
| <b>2018</b> | 0.419                                   | 0.393                 | 0.419                         | 0.416                 |

Source: Author's own calculations. The dataset was obtained from ENGHO 2004/2005, 2012/2013, and 2017/2018.

Table 4 shows that the real Gini coefficient adjusted by the Törnqvist index was 0.01 points higher than the nominal Gini coefficient in 2012. The equivalised household incomes deflated by the QUAIDS-COLI reveal noticeable differences. The comparison documents that income inequality in real terms was 0.034 points higher than nominal income inequality in 2012. Both adjusted Gini coefficients demonstrate that due to the anti-poor behaviour of price movements, real income inequality was higher than nominal inequality.

In 2018, the findings show that income inequality, adjusted by the Törnqvist index, was 0.003 points lower than nominal income inequality. When calculations are undertaken employing the QUAIDS model, the difference is more pronounced at 0.026 points. Both results reveal that Argentina was more equal in real terms in 2018.

## Conclusion

This paper presents a perspective to scholars interested in understanding and measuring the evolution of income inequality. It highlights the importance of considering the impact of price movements and household consumption patterns in examining income inequality trends by showing how the cost of living inequality is reflected in income inequality.

The findings suggest considerable differences in the cost-of-living indices across income groups, especially between 2004 and 2012, which coincided with the 2000s commodity boom and the recovery era. The results show that given the lower-income groups spend a relatively higher fraction of their budget on wage goods, especially on food and beverages, they experienced the highest increase in their cost of living, while the top decile faced the lowest increase. In a nutshell, due to the significant differences in the top and bottom deciles' food budget share

and the sharp rise in food prices, price movements during this period had a regressive impact on income distribution.

Contrary to the commodity boom era, the changes in relative prices were strongly biased against the top decile between 2012 and 2018, seeing as the price indices of luxuries such as leisure and eating outgrew faster than average prices and necessity prices. In addition to these biased price movements, the cut in public utility subsidies contributed to this anti-rich inflation rate. As discussed above, the findings in the literature report that the richest households received approximately twice as much in total energy subsidies compared to the poorest households due to inclusion errors. Therefore, dramatic rises in housing prices and transport & communication can be expected to have made a significant contribution to the cost of living index of the top decile. As the richest households allocated a greater amount of their total expenditure to these consumption bundles than the poor, the sharp rise in the price of transport & communication and housing could explain why the top decile experienced the highest increase in the cost-of-living index during this period.

Comparing nominal and real income inequality indices shows that using the income level-specific cost of living indices provides a different picture of the evolution of income inequality over fourteen years from the one which is estimated by neglecting inflation differentials. From 2004 to 2012, the anti-poor inflation rate translated into an increase in real income inequality that is not captured by standard measures of inequality. In terms of the post-commodity boom period, the findings show that the distribution of income became more equal in real terms after the 2000s commodity boom ended.

Relying on these results, one can say that it is important to allow for the cost-of-living index differentials when measuring income inequality, as neglecting these differences can overestimate or underestimate the changes in income inequality over time. In addition to measuring income inequality, this paper provides an important insight regarding public policymaking. In this paper, group-specific inflation rates are applied to total income, ensuring conceptual consistency with income-side policies such as wage and pension indexation, social security adjustments, and income tax bracket design. While this approach focuses on real income inequality, the results also have direct implications for consumption-side policies—such as targeted VAT reductions or food subsidies—because lower-income households devote most of their budgets to essential goods. Moreover, the findings suggest that the national consumer price index may fall short of representing different subpopulations in Argentina, with potential consequences for both income and consumption policies aimed at promoting inclusive growth.



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The data that support the findings of this study are openly available in Instituto Nacional de Estadística y Censos de Argentina at <https://www.indec.gov.ar/indec/web/Institucional-Indec-BasesDeDatos-4>.

**Conflict of interest:** The author has no conflict of interest to note. Ethical approval is not applicable for this article. Any inaccuracies or oversights remain solely my own responsibility.

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