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Abstract

We investigate the effect that seemingly minor features of the implementation of cost-of-living adjustments have on the distribution of household expenditures, by developing an analytical framework that is consistent with standard consumer theory, and mindful of data limitations faced by practitioners. The main result is at odds with common sense: even when multiple price indices are available (e.g. a food and a non-food CPI), it turns out that using a single price index (e.g. the total CPI), to adjust the consumption aggregate is recommended. The practice of adjusting sub-components of consumption separately (food with a food index, non-food with a non-food index) can lead to a systematic bias in the welfare measure, and consequently in poverty and inequality measures. Using Iran's 2019 Household Expenditures and Income Survey, we find that the bias manifests as a systematic underestimation of urban poverty, and overestimation of rural poverty.

1 Monetary welfare and price adjustment

The measurement of poverty and inequality is grounded in inter-personal welfare comparisons: in order to make broad judgments on social well-being, we need to be able to clearly tell which of any two individuals is "better off". Even when this difficult question is enclosed within the bounds of a welfarist approach (Ravallion, 1994, 2016), and household consumption expenditure is chosen as the basis of the welfare measure, a number of issues need solving before comparisons across individuals can be made soundly. Among the most crucial of these issues is the fact that observed household expenditures typically do not just reflect differences in consumption, and thus welfare, but differences in prices, both over time (inflation) and across areas of a country (geographical cost-of-living differences). Nominal household expenditures need to be adjusted for these differences, to avoid mis-ranking the welfare of individuals facing different prices.

The question of how the adjustment should be performed, exactly, has many theoretical and empirical ramifications, explored by a vast literature (Chakrabarty et al. 2018; Chen et al. 2020; Gaddis 2016; Gibson and Kim 2013, 2015, 2019; Jolliffe 2006; McKelvey 2011; Ray 2018). Much has been written, for instance, on the choice of the index used to deflate the nominal welfare indicator (Deaton and Muellbauer, 1980; Diewert 1983). What concerns economists, and welfare analysts in particular, is not the search for the best measure of price differences, *per se*. Rather, the goal is to find the best approximation of a welfare measure that is consistent with consumer theory. Ideally, this would be accomplished by using a *true cost-of-living index* (TCLI) to adjust nominal expenditure, which maintains the interpretation of the welfare measure as a utility function derived from an expenditure function. However, implementing this approach is usually so challenging empirically, that in most applied work analysts resort to the use of consumer price indices: deflating nominal expenditure by a Paasche index yields an approximation of money-metric utility (MMU) (Deaton and Zaidi, 2002), while using a Laspeyres index leads to an approximation of the welfare ratio (WR) (Blackorby and Donaldson, 1987).[1]

There is a need for guidelines regarding many other practical aspects of price deflation, beyond the choice of an appropriate index, but clear indications are hard to come by. This paper contributes to the literature on price adjustments and welfare measurement by tackling some issues that have so far been left unaddressed: *how many* deflators should be used? A recurring scenario in practice is the

[1] Section 4 of this paper elaborates on this short summary, by defining the objects mentioned here – TCLI, Paasche, Laspeyres, MMU, WR – and going over the implications of the choice of a deflator on the welfare measure.

availability of several price indices with partial coverage – say, food and non-food – alongside an aggregate price index. The paper works out the implications of two alternative price adjustment strategies, namely adjusting the sub-components of household expenditure separately using multiple price indices, versus using a single index to adjust the total.

Findings deliver a few clear-cut suggestions for applied work. While using multiple indices (MI) to adjust the components of nominal household expenditure may seem like the most accurate way to make use of the available information on price variation, we find that whenever we are able to rank the two methods, the single index approach (SI) comes out on top. This seemingly counterintuitive result depends on the fact that our goal is not to adjust nominal expenditures for changes in purchasing power, but rather to derive measures of welfare that are consistently comparable across individuals.

The paper illustrates the impact of the multiple-versus-single index deflation choice on the distribution of real expenditures and on key poverty and inequality estimates, using Iran's 2019 Household Expenditures and Income Survey (HEIS) as a case study. When MI is chosen over SI, poverty is slightly underestimated at the national level, though the most significant differences are observed locally, with a systematic underestimation of poverty in urban areas and, conversely, an overestimation of poverty in rural areas. Arguably, the conditions that generate the results – a wide urban-rural divide in housing costs, with rents being higher in cities, and a comparably narrow spatial variation of food prices – are not specific to Iran. This implies that in countries sharing this feature, a seemingly minor detail in the implementation of spatial adjustments may be cause a systematic and not negligible distortion of the urban-rural poverty profile.

The paper is organized as follows. Section 2 explains the notation used in the rest of the paper; section 3 discusses the implications of the use of multiple indices or a single index to adjust household expenditures; section 4 provides a rationale to rank the two strategies; section 5 provides an empirical illustration of the consequences of said strategies on the estimated distribution of welfare, using Iran's 2019 Household Expenditures and Income Survey (HEIS); section 6 concludes, and offers a summary of the recommended deflation strategy.

2 Price indices: general notation

This section introduces the notation for dealing with price indices, in a way that is convenient for the analysis carried out in the rest of the paper. We denote a generic price index by I – one can think of it as a spatial or a temporal index, indifferently – and will assume that I can be described by a set of three parameters. Suppose that the population consists of H households, with H a positive integer. The first parameter that is needed to define I is $\{r\}$, with $r \leq H$: $\{r\}$ is a partition of the households' set, and defines the *aggregation level* of the price index. If $r = H$, for instance, we say that the index is calculated at the household level; if P is the number of PSUs, and $r = P$, then we say that the index is calculated at the PSU level; if R is the number of regions, and $r = R$ then the index I is defined at the regional level, and so on. If one thinks of a *monthly* temporal CPI, the parameter r would define a partition with $r = 12$. All this can be summarized by denoting the price index as $I(r)$: this makes it explicit than each price index is characterized by a certain aggregation level.

The second parameter needed is $k \leq K$, where K is the total number of goods and services traded and consumed by the households: k is useful to define the *coverage* of the price index I. If $k = K$ we say that the index has full coverage, that is, its elementary components capture each and every commodity consumed by the households. To account for both the aggregation level and the coverage, the generic price index will be denoted as $I(r, k)$.

Before introducing the third parameter, it is convenient to define $L_i(r, k)$ as a Laspeyres price index, where the subscript *i* runs from 1 to r. This is worth illustrating with a couple of examples. Consider a simple *monthly* CPI (e.g., a monthly Laspeyres temporal index): in this case, prices are aggregated on a monthly basis, $r = 12$, and we denote the corresponding twelve monthly Laspeyres indices by $L_i(12, k)$, with $i = 1, 2, \dots, 12$. Similarly, for a spatial price index defined at the urban-rural level, $r = 2$ and the two corresponding indices are denoted by $L_i(2, k)$, with $i = 1, 2$. In sum, the notation $L_i(r, k)$ reminds us that the Laspeyres index a) is defined with a certain aggregation level r, b) it has a certain coverage k, and c) it takes on as many values as indicated by r ($i = 1, \dots, r$). In a similar vein, we can denote a Paasche price index by $P_i(r, k)$. With these two definitions in mind, we introduce a third parameter α and define the following class of price indices:

$$I_i(r, k, \alpha) = [L_i(r, k)]^{\alpha}[P_i(r, k)]^{1-\alpha} \qquad \text{with } \alpha \in \{0, 0.5, 1\} \tag{1}$$

Equation 1 defines a family of three price indices, depending on the value of the parameter α. When $\alpha = 0$, $I_i(r, k, \alpha)$ corresponds to a Paasche price index, when $\alpha = 1$, to a Laspeyres price index, and when $\alpha = 1/2$ to a Fisher price index. In addition, each index is characterized by its own aggregation level r and coverage k. In what follows, $I_i(r, k, \alpha)$ can always be thought of as any of the above three indices, either spatial or temporal, with different aggregation and coverage levels.

A similar notation can be used to define two *sub-indices* restricted to *food* and *non-food* commodities and services. That is useful to deal with situations when prices are only available for food items, or where food- and nonfood price levels vary with different dynamics or gradients, so that the analyst may consider using two deflators instead of one. In order to define the food- and non-food price indices $-\, I_i^f$ and I_i^{nf}, respectively –, we only need to modify equation 1 slightly:

$$I_i^f(r, k_f, \alpha_f) = \left[L_i^f(r, k_f)\right]^{\alpha_f}\left[P_i^f(r, k_f)\right]^{1-\alpha_f} \quad \text{with } \alpha_f \in \{0, 0.5, 1\} \tag{2}$$

$$I_i^{nf}(r, k_{nf}, \alpha_{nf}) = \left[L_i^{nf}(r, k_{nf})\right]^{\alpha_{nf}}\left[P_i^{nf}(r, k_{nf})\right]^{1-\alpha_{nf}} \quad \text{with } \alpha_{nf} \in \{0, 0.5, 1\} \tag{3}$$

where the coverage of the food price index (equation 2) is denoted by $k_f \leq K_f$ (K_f denoting the total number of *food* items traded in the economy), and the coverage of the nonfood index (equation 3) is denoted by $k_{nf} \leq K_{nf}$ (K_{nf} denoting the total number of non-food market commodities and services). Obviously, $K_f + K_{nf} = K$. The parameters α_f and α_{nf} play the same role as α in equation 1: they allow notation to be encompassing of the food Paasche ($\alpha_f = 0$), the food Laspeyres ($\alpha_f = 1$) and Fisher ($\alpha_f = 0.5$) indices, as well of their nonfood counterparts.

As a last step, equations 2 and 3 can be compared with equation 1: the structure is common to all definitions, and allows us to identify a price index depending on whether we need a Laspeyres, a Paasche or a Fisher index (parameters $\alpha, \alpha_f, \alpha_{nf}$), and on the aggregation level (parameter r), and the coverage (parameters k, k_f, and k_{nf}). Table 1 illustrates the use of the seven parameters.

Table 1. A useful parametrization for price indices commonly used in poverty analysis

Type of index	Aggregation level (r)	Coverage (k, k_f, k_{nf})	$(\alpha, \alpha_f, \alpha_{nf})$	Formula
Paasche	$i = 1,..,r$	$j = 1, \dots, k$	$\alpha = 0$	$P_i = \left(\sum_{j=1}^{k} w_j^i \left(\frac{p_j^0}{p_j^i}\right)\right)^{-1}$
Laspeyres	$i = 1,..,r$	$j = 1, \dots, k$	$\alpha = 1$	$L_i = \sum_{j=1}^{k} w_j^0 \left(\frac{p_j^i}{p_j^0}\right)$
Fisher	$i = 1,..,r$	$j = 1, \dots, k$	$\alpha = 0.5$	$F_i = (L_i \times P_i)^{1/2}$
Food Paasche	$i = 1,..,r$	$j = 1, \dots, k_f$	$\alpha_f = 0$	$P_i^f = \left(\sum_{j=1}^{k_f} w_j^i \left(\frac{p_j^0}{p_j^i}\right)\right)^{-1}$
Food Laspeyres	$i = 1,..,r$	$j = 1, \dots, k_f$	$\alpha_f = 1$	$L_i^f = \sum_{j=1}^{k_f} w_j^0 \left(\frac{p_j^i}{p_j^0}\right)$
Food Fisher	$i = 1,..,r$	$j = 1, \dots, k_f$	$\alpha_f = 0.5$	$F_i^f = \left(L_i^f \times P_i^f\right)^{1/2}$
Nonfood Paasche	$i = 1,..,r$	$j = 1, \dots, k_{nf}$	$\alpha_{nf} = 0$	$P_i^{nf} = \left(\sum_{j=1}^{k_{nf}} w_j^i \left(\frac{p_j^0}{p_j^i}\right)\right)^{-1}$
Nonfood Laspeyres	$i = 1,..,r$	$j = 1, \dots, k_{nf}$	$\alpha_{nf} = 1$	$L_i^{nf} = \sum_{j=1}^{k_{nf}} w_j^0 \left(\frac{p_j^i}{p_j^0}\right)$
Nonfood Fisher	$i = 1,..,r$	$j = 1, \dots, k_{nf}$	$\alpha_{nf} = 0.5$	$F_i^{nf} = \left(L_i^{nf} \times P_i^{nf}\right)^{1/2}$

Note: the table shows the formulas for the different parametrizations illustrated in equations 1, 2 and 3.

To sum up, the notation introduced in this section identifies a set of seven parameters $(r, k, k_f, k_{nf}, \alpha, \alpha_f, \alpha_{nf})$, that are used in the rest of the paper to produce theoretical propositions on how to carry out both spatial and temporal price adjustments.

3 Single versus multiple price indexes in welfare deflation

An important choice faced by practitioners pertains to the use of a single price index versus the use of multiple (two or more) sub-indexes. The simplest case is the use of two different price indexes, for instance a food price index to deflate the food component of the consumption aggregate and a non-food price index to deflate the non-food component. In what follows, we develop a theoretical framework that shows how the use of these two alternative approaches leads to different welfare distributions, and has systematic implications in terms of welfare *ranking* of households.[2]

Following the notation introduced in section 2, consider the following two alternative price adjustment procedures. A first method of deflation uses a single index (SI):

$$x_{SI}^{h} = \frac{x^{h}}{I_i(r,k,\alpha)} \tag{4}$$

where x^h is the total nominal expenditure of household *h* belonging to area (or period) *i*, and the index $I_i(r,k,\alpha)$ is defined in equation 1.

A second method consists in using a multiple index (MI) adjustment method:

$$x_{MI}^{h} = \frac{x_f^{h}}{I_i^{f}(r,k_f,\alpha_f)} + \frac{x_{nf}^{h}}{I_i^{nf}(r,k_{nf},\alpha_{nf})} \tag{5}$$

where x_f^h is the total nominal expenditure on food items of household *h* belonging to area (or period) *i*, and x_{nf}^h is the nominal expenditure on all the good and services included in the consumption aggregate but different from food.

The first result identifies conditions under which the use of single versus multiple price indexes leads to the same results in terms of welfare ordering and distribution.

Proposition 1 (Invariance).

Assume $(r,k,k_f,k_{nf},\alpha,\alpha_f,\alpha_{nf}) = (H,K,K_f,K_{nf},0,0,0)$, then $x_{SI}^h = x_{MI}^h$.

Proof: see the appendix.

Proposition 1 establishes that when using a *household-level* Paasche index, one inclusive of all items (food and nonfood), then using a single or multiple index approach leads to the same real consumption aggregate: $x_{SI}^h = x_{MI}^h$. In other words, the choice of the method is irrelevant for the estimation of poverty and inequality indices, as well as for any other moment of the empirical distribution function.

[2] While in principle the multiple index approach can be implemented with more than two subindices, each matched with the appropriate component of the consumption aggregate, we focus on the use of *two* subindices (food and nonfood) to keep the tractability of the algebra within reasonable limits. Arguably, the food-nonfood split is also most common in applied work.

The situation described by Proposition 1 corresponds to the conditions assumed by Deaton and Zaidi (2002).

Does such a Paasche index exist in practical work? Arguably, it's unlikely for two main reasons. First, household survey data that are typically used for welfare measurement do not include information on prices faced by each household. Rather, households within a defined geographical area are associated with same set of prices, either because expenditure data collected at household level are matched with prices from a common market in the area or because prices are proxied by unit values.[3] In either case, the lack of household level price data leads to $r \neq H$, and therefore $x_{SI} \neq x_{MI}$. Second, the ideal scenario of full coverage of prices for all items included in the consumption aggregate (i.e. $k = K$) is rarely realized, either because not all items are sold in the market of reference (in case household questionnaires are complemented by market level questionnaires) or because is often impossible to collect information on quantities consumed for non-food items and hence is not possible to proxy prices with unit values.

By relaxing the assumptions of Proposition 1, we obtain a new theoretical result, that clarifies how important is, in practice, the choice between the use of one deflator (SI) versus two deflators (MI).

Proposition 2 (Distributional bias).

Assume $(r, k, k_f, k_{nf}, \alpha, \alpha_f, \alpha_{nf}) = (r, k, k_f, k_{nf}, \alpha, \alpha, \alpha)$, and $k < K$ or $r < H$ or both, then $x_{SI}^h \neq x_{MI}^h$. Moreover, if $I_i^f > (<) I_i^{nf}$, then $x_{SI}^h - x_{MI}^h > 0$ if and only if:

$$w_h > (<)\, \Delta_i = \frac{\left(I_i - I_i^{nf}\right)}{\left(I_i^f - I_i^{nf}\right)} \frac{I_i^f}{I_i}$$

where w_h is the food budget share of household *h*, and Δ_i is always strictly positive.

Proof: see the appendix.

Proposition 2 provides a clear-cut result. In a nutshell, the proposition says that as soon the "ideal" Paasche index (one covering *all* commodities in the economy, and calculated at for each household in the sample) is not available, and therefore the analyst commits to a second-best solution, then the choice of method SI versus MI implies a *systematic* difference between the *real* consumption

[3] As discussed in DZ, the most common approach when using unit values is to replace the individual p_k^h by their medians over households in the same PSU or locality in order to minimize noise and address potential problems related to outliers (DZ, p. 42).

aggregates computed using the two approaches: $x_{SI}^h \neq x_{MI}^h$. In particular, the difference between the two approaches depends on two factors. First, on the "price environment" in which the household lives: the sign of the difference between x_{SI}^h and x_{MI}^h depend on whether the food price index is greater or smaller than the nonfood price index. Second, the household consumption pattern: the sign of the difference between x_{SI}^h and x_{MI}^h depends on how large the household's food budget share w_j^h is. When $I_j^f > I_j^{nf}$ and w_j^h is above a certain threshold (Δ_i in proposition 2), then using multiple indexes implies $x_{MI}^h < x_{SI}^h$.

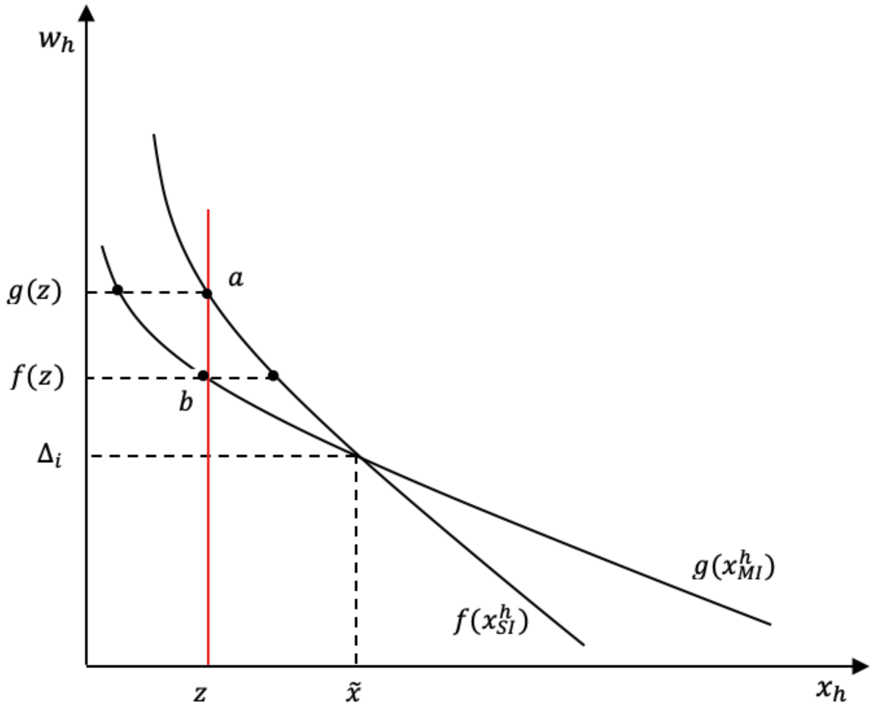
It is worth pausing and elaborating on the fact that MI implies a systematic "bias" with respect to SI: in any economic environment where food prices are higher than non-food prices, the use of MI will be responsible for lower levels of standard of living for households with a relatively high food budget share, and x_{MI}^h will be lower than x_{SI}^h the higher is the importance of food in the overall household budget. With respect to SI, the MI approach will tend to penalize households who spend a large share of their budget in the category of goods that is relatively more expensive. How much is "a large share"? The threshold is defined on food budget shares. So, when food is relatively more expensive ($I_i^f > I_i^{nf}$), households with "high" food budget shares will be penalized by MI. Conversely, when non-food is relatively more expensive ($I_i^f < I_i^{nf}$), households with "low" food budget shares will be penalized by MI. How high or how low is identified precisely by Δ_i.

The last statement can be illustrated and interpreted more meaningfully for welfare analysis if we assume that the Engel law applies, i.e. that poorer households devote a higher budget share on food compared to non-food. In this case the use of multiple index deflation implies welfare measures *systematically lower* than estimates based on the single index approach, and the size of the gap tends to be larger the poorer is the household. Hence, proposition 2 tells us that, if $I_i^f > I_i^{nf}$, the using multiple indexes tends to make poorer households poorer with respect to using a single index.[4] Figure 1 makes explicit use of Engel curves, which allows to project the food budget shares (vertical axis) on the metric of *real* expenditures (horizontal axis) and vice-versa. Depending on whether the analyst uses SI or MI approach, there are two possible Engel curves: $w_h = f(x_{SI}^h)$ and $w_h = g(x_{MI}^h)$ with

[4] This is what happens only if the threshold value Δ_i is not too high: when Δ_i is "very" high, then the condition $w_h > \Delta_i$ will not be satisfied, neither for poor nor for nonpoor households and using multiple index deflation will "reward" all households and lead to consistently lower FGT estimates compared to a single deflator approach.

$f(\cdot)$ and $g(\cdot)$ real continuous function such that $f' < 0$ and $g' < 0$. Assume that a value $\Delta_i \in (0,1)$ exists such that $f^{-1}(\Delta_i) = g^{-1}(\Delta_i) = \tilde{x} > 0$. We know, from proposition 2, that such a value is unique, which means that the functions $f(\cdot)$ and $g(\cdot)$ intersect only once. In particular, if $I_i^f > (<) I_i^{nf}$ and $w_h > \Delta_i$, then $x_{SI}^h - x_{MI}^h > (<) 0$ and vice-versa for $w_h < \Delta_i$. Figure 1 illustrates the case for $I_i^f > I_i^{nf}$.

Figure 1. Engel's law and price adjustment ($I_i^f > I_i^{nf}$)



Source: Authors' elaboration.

The advantage of "projecting" the food budget shares on expenditures by means of the Engel curves is that we can compare budget shares with the poverty line z.[5] Suppose, for instance, that $z < \tilde{x}$, as in Figure 1. Then, all the households such that $f(z) < w_h < g(z)$ will be classified as poor if we use multiple indexes (the Engel curve lies to the left of the poverty line) to adjust the real consumption

[5] We assume the poverty line to be independent from price deflation. Poverty analysts always have one of two options: either use a nominal poverty line with a price-adjusted consumption aggregate, or a price-adjusted poverty line (i.e., multiple poverty lines) with a nominal consumption aggregate. Figure 1 uses option 1.

aggregate and will be classified as non-poor if we use a single index (the Engel curve lies to the right of the poverty line). The conclusion would reverse when $I_i^f < I_i^{nf}$. In any case, depending on the relative values of the food and non-food price indices and on the position of poverty line with respect to $\tilde{x}$, thanks to the Engel Law, proposition 2 makes it possible to identify (profile) the households that would change their poverty status according to the different deflation procedures. All households lying on the segment $a - b$, in Figure 1 would be classified as non-poor if SI were used, while will be classified as poor if MI were used.

Finally, it is interesting to note that if $z = \tilde{x}$ no household would change its poverty status, which implies that the headcount poverty rate would be insensitive to the deflation procedure, while higher-order FGT-poverty measures and other poverty measures that are sensitive to the distance from the poverty line would be affected by the choice between the SI and MI deflation.

The general result enunciated in proposition 2 is easier to interpret when we focus on the Laspeyres index ($\alpha = 0$), as shown in the following Corollary.

Corollary (Distributional bias with Laspeyres index)

Assume $(r, k, k_f, k_{nf}, \alpha, \alpha_f, \alpha_{nf}) = (r, k, k_f, k_{nf}, 1,1,1)$, and $k < K$. If $L_i^f > (<) L_i^{nf}$, then $x_{SI}^h - x_{MI}^h > 0$ if and only if:

$$w_h > (<) \Delta_i = w_i \frac{L_i^f}{L_i}$$

where w_i is the average food budget share in region (or period) i.

Proof: see the appendix.

When both single and multiple price index deflation approaches use a Laspeyres index, the threshold value Δ_i is given by the food budget share w_i, determined by the aggregation level of the index (for example, w_i may be the average food budget share in region *i*, and/or time period *i*), multiplied by L_i^f / L_i, that is the ratio of the food- to the general price index. If $L_i^f > L_i^{nf}$, the ratio is always greater than 1 and the condition $w_h > \Delta_i$ is satisfied by households with a food budget share strictly larger than w_i. If the price indices have been computed by using the food budget shares of the poor households[6], then, because of the Engel law, $w_i \cong max\{w_h\}$ and therefore the condition $w_h > \Delta_i$ can be hardly binding. This implies that the use of the MI approach increases households' consumption

[6] On this, see Deaton and Zaidi (2002).

aggregates with respect to the SI approach for *almost all* households: $x_{MI}^{h} > x_{SI}^{h}$ for most h. The wedge in the consumption aggregates is such that all poverty counts and gaps are expected to be affected, as well as inequality indices with concave evaluation functions (GEIs, Atkinson's indices with a inequality aversion parameter strictly greater than 1, etc.).

4 Which approach is better?

Results presented in section 3 identify the existence of systematic biases related to the choice of a single versus multiple index deflation approach, but they do not have immediate implications from the *normative* viewpoint. Is it possible to express a preference for either approach to adjust the nominal consumption aggregate? Is the SI approach "better" than the MI approach? To answer this question, we need to identify a benchmark, a reference theoretical measure of welfare that we should approximate by using the SI or the MI adjustment procedure.

This section is technical in its content and can be skipped by readers who are not interested in the process that leads to our result. The discussion is organized into two separate blocks. Section 4.1 reviews the theory underlying the results obtained in section 4.2.

4.1 The theoretical framework

The *money metric utility function* (MMU) is a convenient monetary measure of individual welfare (Samuelson 1974). The MMU^h function is defined as the *minimum expenditure* for reaching the utility level $u(q^h)$ associated to the consumption bundle q^h, given the market price vector p^h:

$$MMU^h = c(u(q^h), p^h) \qquad (6)$$

where $c(u(q^h), p^h)$ denotes the *expenditure* or *cost function* for household h (Deaton and Muellbauer 1980: 47, McKenzie 1957).

Under the standard assumptions that households are rational and minimize their expenditure, the empirical counterpart of MMU^h in equation 6 is given by $p^h q^h$, where q^h is the consumption bundle that, given prices p^h, provides the utility $u(q^h)$ at the minimum cost. The household expenditure $p^h q^h$, however, would not be comparable across households: both equation 6 and $p^h q^h$ cannot be used to carry out interpersonal or intertemporal welfare comparisons.[7] The problem can be overcome by introducing a *reference price vector* p^0. Hence, instead of equation 6, we can consider the following money metric utility:

$$MMU^h = c\left(u\left(q^h\right), p^0\right) \tag{7}$$

Equation 7 would be both utility consistent and comparable across households, but it suffers from a major drawback: preferences cannot be observed and equation 7 cannot be estimated based on the consumption behavior of household h available from survey data. Household h minimizes its expenditure given the market price vector p^h, not the reference price vector p^0.

A possible solution is to introduce the *Paasche-Konüs true cost of living index* ($TCLI_{PK}^h$), defined as follows:

$$TCLI_{PK}^h = \frac{c\left(u\left(q^h\right), p^h\right)}{c\left(u\left(q^h\right), p^0\right)} \tag{8}$$

The true-cost-of-living index (TCLI) defined in equation 8, originally due to Konüs (1939), has been used widely in applied economics (Ray 2018). Interpretation is as follows: for household h, $TCLI_{PK}^h$ compares the cost of reaching a fixed level of utility $u\left(q^h\right)$ at two different price situations p^h and p^0, such as the prices in two time periods or those in different locations.

Given equation 8, equation 7 can be rewritten as follows:

$$MMU^h = \frac{c\left(u\left(q^h\right), p^h\right)}{TCLI_{PK}^h} = \frac{p^h q^h}{TCLI_{PK}^h} = \frac{x^h}{TCLI_{PK}^h} \tag{9}$$

[7] To see why, suppose that household *h* and household *j* face different market prices, say p^h and p^j, respectively; suppose that household *j* achieves the same utility level of household *h*, $u(q^h)$, but does so with different expenditure level, due to different prices it faces:

$$c(u(q^h), p^h) \neq c\left(u(q^h), p^j\right)$$

When this is the case, then MMU^h and MMU^j differ, even if the utility level is the same for the two households: the use of MMU as a welfare measure would lead to misranking households *h* and *j* (when two households achieve the same utility level they should be ranked equally well off) where x^h is the total expenditure of household *h*. If the analyst is able to estimate $TCLI_{PK}^h$, the expression in equation 9 would be *i)* utility-consistent, *ii)* observable, and *iii)* suitable for interpersonal comparisons.

The problem with $TCLI_{PK}^h$ is precisely its estimation, which requires the specification of a utility function, and the estimation of the preference parameters for its application. Deaton and Zaidi (2002) – henceforth DZ – put the issue as follows: "The exact calculation of money metric utility requires knowledge of preferences. Although preferences can be recovered from knowledge of demand functions, we typically prefer some shortcut method that, even if approximate, does not require the estimation of behavioural relationships with all the accompanying assumptions, including often controversial identifying assumptions, and potential loss of credibility." For this reason it can be argued that equation (4) does not qualify as a computationally convenient welfare measure.

The shortcut method mentioned by DZ consists in approximating the MMU^h in equation 9. The idea is simple and ingenious: use a first order Taylor expansion of MMU^h to show that:

$$MMU^h \approx \frac{x^h}{P^h} \tag{10}$$

where $P^h = p^h q^h / p^0 q^h$ is a household-level Paasche index. MMU^h in equation 10, corresponding to equation 2.6 in DZ's original paper: that is a monetary welfare measure compliant with a number of desirable properties: it can be calculated by the ratio of total nominal household expenditure (x^h) to a household-level Paasche index (P^h). The relation is now an approximation, whose precision ultimately depends on the gap between the prices faced by the household and the reference set of prices. If we compare equation 9 and equation 10, it is useful to note that the Paasche index P^h can be interpreted as a proxy for the $TCLI_{PK}^h$: the closer is P^h, empirically, to $TCLI_{PK}^h$, the better DZ's shortcut in equation 10 is expected to work.

Blackorby and Donaldson (1985, 1987, 1988) – BD henceforth – observed that MMU^h is not necessarily a concave function of x^h.[8] This creates problems in evaluating redistributive policies, as it may lead to results that do not satisfy the principle of transfer: the use of MMU, in other words, comes with the risk of using distributionally insensitive tools, such as poverty and inequality measures

[8] The reason MMU in equations 9 and 10 is not concave is that $TCLI_{PK}^h$, at the denominator, depends on the utility level u, whose functional form is not necessarily known.

obtained by aggregating MMU (Foster, Greer and Thorbecke, 1984). To overcome this issue, BD proposed an alternative welfare measure, the *welfare ratio*:

$$WR^h = \frac{c(u, p^h)}{c(u_z, p^h)} \tag{11}$$

The welfare ratio is the ratio between the expenditure function of household h and the minimum expenditure necessary to achieve a reference utility level u_z (typically, the utility level around a poverty line *z*). In equation 11, the expenditure $c(u_z, p^h)$ is a sort of numeraire, so that the WR^h can be interpreted as a multiplier of the minimum expenditure necessary to achieve u_z. If $WR^h = 1.2$, for instance, then the current expenditure for household *h* is 20% higher than the expenditure needed to achieve u_z. If the expenditure at the denominator is interpreted as a poverty line, then WR^h tells how many poverty baskets can be purchased by the current spending of household *h*.
To make WR^h operational, it needs to be transformed into a monetary value, which can be done simply by multiplying WR^h by $c(u_z, p^0)$, the monetary cost of u_z at a reference price vector p^0:

$$\widetilde{WR}^h = \frac{c(u, p^h)}{c(u_z, p^h)} c(u_z, p^0) \tag{12}$$

Equation (7) can be simplified by introducing the *Laspeyres-Konüs true cost of living index*:

$$TCLI_{LK}^h = \frac{c(u_z, p^h)}{c(u_z, p^0)} \tag{13}$$

The TCLI in equation 13 can be compared with the Paasche-Konüs index in equation 8. Conceptually, they are both TCLIs, and they both serve the purpose of adjusting nominal expenditures for differences in purchasing power. The difference between the two indices is the reference utility level used in their definitions: in equation 8 Paasche uses u^h, while in equation 13 Laspeyres uses u_z.
By plugging-in equation (8) in equation (7), we obtain the following result:

$$\widetilde{WR}^h = \frac{c(u, p^h)}{TCLI_{LK}^h} = \frac{x^h}{TCLI_{LK}^h} \tag{14}$$

To interpret equation 14 it is convenient to compare it with equation 9. The key difference is that in the former $TCLI_{LK}^{h}$ does not depend on household expenditure x^{h}, so that $\widetilde{WR}^{h}$ is a linear function of x^{h}.

Regarding equation 14, DZ observe that to the extent to which is plausible to assume that the consumption bundle around the poverty line is insensitive to market prices, then $TCLI_{LK}^{h}$ can be approximated by a *Laspeyres price index*:

$$\widetilde{WR}^{h} \approx \frac{x^{h}}{L^{h}} \tag{15}$$

Both the welfare ratio WR defined in equation 15, and the money metric utility MMU defined in equation 10, satisfy the properties identified at the beginning of this section, and therefore represent strong candidates for proxying the living standards based on household budget survey data.

The common structure of the two measures facilitates their interpretation and practical implementation: they are both defined as the ratio between nominal total household expenditure to a true-cost-of-living index. They can both be approximated by replacing TCLI with a price index, a Paasche index in the case of MMU, a Laspeyres index in the case of the WR. The choice of the index (whether Paasche or Fisher) implies a different choice of the underlying price measure. In this sense, the question "which index is the best choice" is not a good question: No 'best' index exist. By choosing the price index, the analyst chooses the welfare measure: Paasche implies a MMU, Laspeyres implies a WR, Fisher, Törnquist and other indices imply different welfare measures.

The analysis reviewed in this section hopefully helps to clarify the recommendation put forward in Deaton and Zaidi (2002) and confirmed by Mancini and Vecchi (2022), and sets the stage of the analysis carried out the next section. The aim is not to contribute the debate on the best welfare measure (MMU versus WR), but rather to focus on different price adjustment procedures, SI versus MI, *taking as given that the appropriate welfare concept has been chosen*.

4.2 Can the SI and the MI methods be ranked?

Section 4.1 discusses two welfare concepts, *MMU* and *WR*, which we use here as benchmarks for assessing whether the practice of using one deflator (SI) is preferable than the practice of using two deflators (MI). Consider first the MMU approach. According to equation 9, the money-metric utility function for household *h*, MMU^{h}, can be written as the ratio between household *h* expenditure and the true-cost-of-living index (TCLI) associated to the utility level *u* achieved by household h at the prevailing prices p^h:

$$MMU^h = \frac{x^h}{TCLI_{PK}^h}$$

where $\mathrm{TCLI}_{\mathrm{PK}}^{\mathrm{h}}$ is the Paasche-Konüs TCLI introduced in equation 8. The equation above, corresponding to equation 9, identifies our benchmark. While methods SI and MI can be assessed against this theoretical benchmark, in practice this is difficult to achieve as the estimation of $TCLI_{PK}^h$ requires knowledge of household preferences, which are generally unobserved: comparing the "first best" with the SI and MI approach is an almost impossible *empirical* exercise. In this section we draw some *theoretical* conclusions by imposing mild restrictions on household preferences.

Proposition 3 (SI method is preferable to MI method when MMU is the welfare measure)
Assume that household h's preferences ($\succcurlyeq_h$) can be represented by a continuous, increasing, and concave utility function, and that $\left(H, k, k_f, k_{nf}, \alpha, \alpha_f, \alpha_{nf}\right) = \left(H, k, k_f, k_{nf}, 0,0,0\right)$. Then, if $I_i^f > (<) I_i^{nf}$ and $w_h < (>) \Delta_i$, then SI approach provides a better estimate of MMU^h than the MI approach.
Proof: see the appendix.

The result illustrated in proposition 3 depends on the *ordering* of $TCLI_{PK}^h$ and the Paasche price index. As discussed in Section 4.1, if preferences are regular, we know that $TCLI_{PK}^h > P_i(H,k)$ which implies $MMU^h < x_{SI}^h$. In this case, whenever conditions determined by the price environment and household consumption patterns determine that MI approach tends to overestimate the real expenditures with respect to SI approach, i.e. if $I_i^f > I_i^{nf}$ and $w_h < \Delta_i$, using the MI approach produces an even larger deviation from MMU^h compared to the SI approach. It should be observed that we cannot achieve a symmetric conclusion if $x_{MI}^h < x_{SI}^h$. In this last case we could have $x_{MI}^h < MMU^h < x_{SI}^h$ and the bias induced by the MI approach can be, in absolute terms, either larger or smaller than the bias implied by the SI approach. However, while – as shown by proposition 3 - we can identify conditions under which SI approach leads to better approximation of MMU^h compared to the MI approach, we cannot identify any conditions under which MI should be preferred to SI.

In line with what discussed in section 4.1, a similar analysis can be carried out using the welfare ratio measure, as defined in equation 14 and here repeated for convenience:

$$\widetilde{WR}^h = \frac{c(u, p^h)}{TCLI_{LK}^h}$$

Proposition 4 (SI method is preferable to MI method when WR is the welfare measure)
Assume that household h preferences ($\succcurlyeq_h$) can be represented by a continuous, increasing, and concave utility function and that $(H, k, k_f, k_{nf}, \alpha, \alpha_f, \alpha_{nf}) = (H, k, k_f, k_{nf}, 1,1,1)$. Then, if $I_i^f >$ $(<) I_i^{nf}$ and $w_h > (<) \Delta_i$ SI approach provides a better estimate of $\widetilde{WR}^h$ than MI approach.
Proof: see the appendix.

As for the case of MMU, proposition 4 shows that we cannot identify conditions that would lead to conclude unambiguously in favor of a MI approach.
Overall, propositions 3 and 4 show that, irrespective of choice of the welfare measure, the SI approach weakly dominates the MI approach. The result is somewhat surprising, if seen from the angle statisticians are used to: the use of multiple deflators does make sense when one seeks consistency between numerators (say, food and nonfood expenditures) and denominators (the deflators). Clearly, the use of a food deflator and a nonfood deflator does a better job in capturing the change in purchasing power of the two sub-aggregates. From the economist's angle things look different. The main aim of welfare analyst is to approximate a well-being measure that allows to make consistent interpersonal comparisons. From this point of view a single deflator could be a preferable solution.

5 An empirical illustration

In this section, we illustrate the empirical implications of choosing a single index (SI) versus multiple index (MI) approach in spatial (and temporal) price adjustment, using data from Iran's 2019 Household Expenditures and Income Survey (HEIS).

5.1 Construction of single and multiple price indexes

Similar to most expenditure surveys, the Iran's HEIS only collects information on expenditure and quantities consumed for food items. Unit values obtained from reported food expenditures and quantities allow for the estimation of a survey-based food spatial price index. On the other hand, when it comes to the measurement of non-food spatial price variation, options are characteristically limited. One of these options is to focus on rent expenditures[9], pin down the characteristics of a "typical" dwelling and predict what the average rent for this dwelling would be in different areas of the country, via hedonic regression. Ratios between predicted prices for each region and the country average can be interpreted as a (Paasche) price index constructed using only one commodity.

The question examined in this section is how to proceed in the construction of single and multiple indexes that are consistent with the theoretical framework presented in previous sections, *after* the food and rent deflators have been estimated. The approach followed in this empirical application for the construction of a "single" index and corresponding food and non-food price indexes are summarized in the equations below:

$$I_i = \left[\sum_{j=1}^{k_f} w_j^i \frac{p_j^0}{p_j^i} + w_{nf}^i \frac{p_r^0}{p_r^h}\right]$$

$$I_i^f = \left[\sum_{j=1}^{k_f} w_j^i \frac{p_j^0}{p_j^i}\right]$$

$$I_i^{nf} = \left[\frac{p_r^0}{p_r^h}\right]$$

where j denotes each food item (belonging to the food commodity group), w_{nf}^i denotes the budget share of non-food expenditure (including rent, and all other non-food items), p_r denotes predicted rents. Notice that in this application, we use the Paasche formula to estimate the single spatial price index (equivalently, $\alpha = 0$). As in section 3, $i = 1, \ldots, r$, where r refers to the survey strata defined at the region – urban/rural level.

Correspondingly, the formulas used for the construction of the spatially adjusted aggregate using SI and MI approach are the following:

$$x_{SI} = \frac{x}{\left[\sum_{j=1}^{k_f} w_j^i \frac{p_j^0}{p_j^i} + w_{nf}^i \frac{p_r^0}{p_r^h}\right]} \tag{18}$$

[9] Rent expenditures constitutes of actual rents for tenants and self-reported rents for owners/non-market tenants. Hedonic regressions have been used to validate self-reported rent estimates.

$$x_{MI} = \frac{x_f}{\left[\sum_{j=1}^{k_f} w_j^i \frac{p_j^0}{p_j^i}\right]} + \frac{x_{nf}}{\left[\frac{p_r^0}{p_r^h}\right]} \tag{19}$$

In this application, the non-food deflator is a single commodity index (based on rent only): this is tantamount to assuming that the spatial price dynamics of all non-food items is the same as the one captured by housing prices. Reflecting the inherent non-tradable nature of housing services and typical tradeable nature of food items, this deflation approach is expected to maximize the relative spatial variability of the non-food price index compared to the food price index, and it is hence well suited to showcase empirical implications of proposition 2.[10]

Regarding the temporal dimension of price adjustments, the *aggregate* monthly CPI index, as estimated by the Statistical Center of Iran (SCI), is used throughout the analysis[11].

5.2 Results

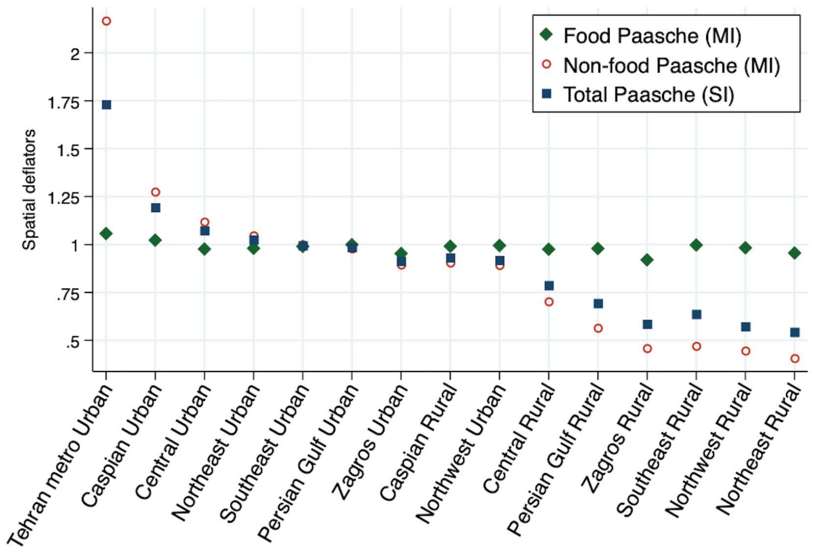
Proposition 2 delivers a prediction as to the sign of the difference between a welfare aggregate deflated using a single index and one deflated using a multiple index approach. According to proposition 2, this difference would depend on the relative level of food and non-food price indexes as well as on observed consumption patterns (implicit differences in the level of welfare), i.e. on the food share level. What makes the analysis of data for Iran particularly useful for an empirical illustration of Proposition 2 is the significant spatial variation observed in Iran both in terms of food shares as well as in terms of food and non-food price indexes (Figure 2).[12]

Figure 1. Food, non-food, and total price indexes

[10] In practice, this assumption may seem too extreme, and welfare analysists may be reluctant to adopt it. A possible alternative, feasible under typical data constraints, would be to assume that the prices of non-food/non-rent items do not vary spatially, and that only housing expenditures are associated to the price dynamics of rent. Results obtained using this alternative approach are available in Appendix 2.

[11] This is to maintain a clean connection with the theory presented in section 3. Results capture entirely and exclusively the impact of the SI and MI approaches for *spatial* deflation.

[12] Another advantage of considering the Iran case is the existence of a well-developed housing / rental market in both urban and rural areas, which allows for an accurate estimation of the rent deflator.



According to Figure 2, the variation of the non-food price index (I_i^{nf}) across regions is much higher compared to that of the food price index (I_i^{f}), reflecting the non-tradeable vs tradeable nature of the items included in each index's basket. Moreover, $I_i^f < I_i^{nf}$ in urban areas, whereas the opposite ($I_i^f > I_i^{nf}$) is observed in rural areas (mostly due to the dynamics of housing prices).

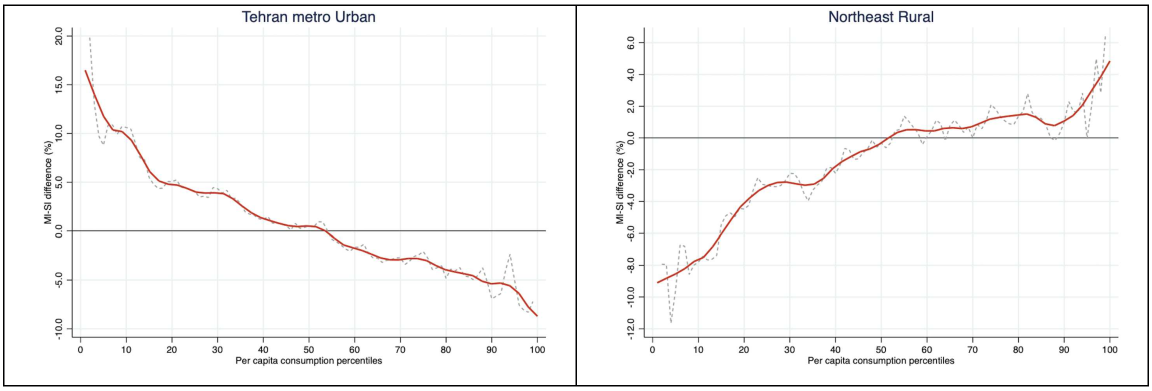
In the rest of the section, we assess the distributional bias of using the MI versus the SI approach in the regions at the two ends of the spectrum pictured in Figure 1: urban Tehran and Northeast Rural. Table 2 displays all the pieces of information entering proposition 2: the top panel reports the values of the food and non-food spatial price indices in each region, as well as the value Δ, calculated as in proposition 2. Δ is a threshold that determines which households (those with food expenditure shares higher or lower than Δ) will see their real expenditures decrease when switching from SI to MI. In particular, in Teheran the food price index is smaller than the non-food price index, and Δ is equal to 24; therefore, proposition 2 implies that switching from a SI deflation strategy to a MI strategy will *decrease* real expenditures of all households with a food budget share *lower* that 24%. Vice-versa, in the rural Northeast, the food price index is larger than the non-food price index, and Δ is equal to 43.7; accordingly, switching from a SI deflation strategy to a MI strategy will *decrease* real expenditures for all households with a food budget share *higher* than 43.7%. The distributional effects of the two scenarios are clarified by the bottom panel of Table 2, which shows average budget shares by decile of (nominal) per capita expenditure. According to proposition 2, we should expect the MI approach to "underestimate" welfare compared to the SI approach for the top 4 deciles of the welfare distribution in urban Tehran, and for the bottom 7 deciles in rural Northeast. In Table 2, deciles where households are "penalized" by MI have been highlighted by using the bold type.

Table 2. Implications of Proposition 2

	Teheran metro urban	Northeast rural
Food Paasche (I^f)	1.06	0.96
Non-food Paasche (I^{nf})	2.17	0.41
Δ	24.0	43.7
Proposition 2 implies:	$x_{MI}^h < x_{SI}^h$ when $\boldsymbol{w_h} < \Delta$	$x_{MI}^h < x_{SI}^h$ when $\boldsymbol{w_h} > \Delta$
Per capita expenditure deciles	**Average food budget shares, w_h (%)**	
1	40.8	**53.0**
2	33.0	**49.5**
3	28.9	**47.5**
4	27.8	**44.6**
5	25.3	**44.3**
6	23.3	**42.1**
7	**20.8**	**42.9**
8	**20.2**	40.9
9	**18.1**	42.7
10	**16.1**	41.6

Note: Expenditure deciles where $x_{MI}^h < x_{SI}^h$ are in bold type.

Figure 2. Percent difference in consumption aggregate deflated using MI and SI approaches, by percentile of the distribution in urban Tehran and rural Northeast



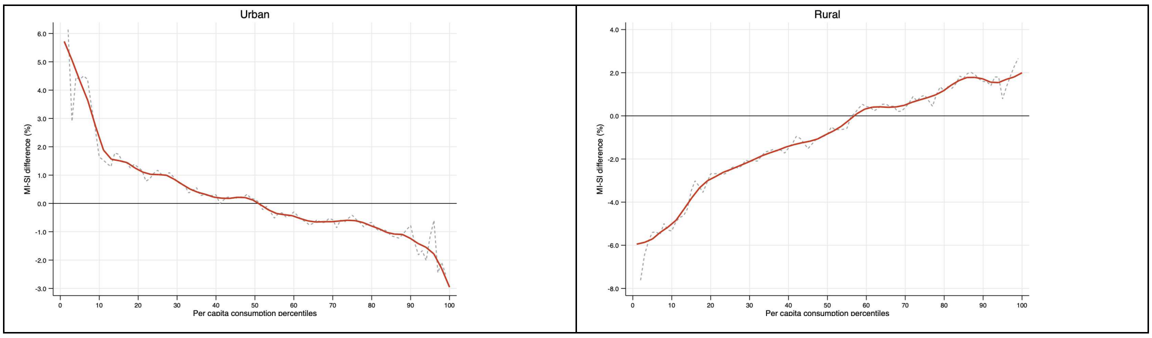
Note: The grey dashed line represents percentage differences at each percentile, while the solid red line is obtained via Kernel-weighted local polynomial smoothing.

Theoretical predictions illustrated in Table 2 are confirmed when looking at the empirical distribution of the bias between the consumption aggregate deflated using the MI approach and the one deflated using the SI approach. Figure 2 is constructed in a way similar to the growth incidence curve (GIC) introduced in Ravallion and Chen (2003): the curve is calculated by computing average real per capita expenditures within each of 100 percentile groups, separately by deflation approach (SI versus MI). These averages by percentile groups are then compared with each other, and a percent difference is computed.[13]. More than as growth incidence curves, then, these graphs can be interpreted as "deflation incidence curves", showing the impact on the welfare measure of switching from SI to MI.

In line with predictions of proposition 2, the results in Figure 2 indicate that – compared to a single index approach (equation 18) – a multiple index approach (equation 19) tends to "under-value" consumption for relatively poorer households in rural Southeast, whereas it tends to "over-value" consumption for relatively poorer households in urban Tehran. Moreover, the magnitude of the bias between a welfare aggregate deflated using a MI approach and one deflated using SI approach varies at different points of the welfare distribution (i.e. as the food share varies). A similar pattern emerges when considering Iran's urban areas as a whole: deflating welfare using the MI approach tends to "over-value" consumption for relatively poorer households in urban areas (Figure 3).

[13] For example, in Teheran, within the first percentile group, average real expenditure when using SI is about 738 thousand Rial/person/month, and 899 thousand when using MI. The difference is 21% of what we consider the "initial value", SI.

Figure 3: Percent difference in consumption aggregate deflated using MI and SI approaches, by percentile of the distribution in urban and rural areas



Note: The grey dashed line represents percentage differences at each percentile, while the solid red line is obtained via Kernel-weighted local polynomial smoothing.

As discussed in section 4, systematic differences between SI and MI approach in welfare deflation have clear implications in terms of poverty measurement, with the magnitude of the bias being sensitive to the position of the poverty line along the welfare distribution, as well as to the poverty measure adopted in the analysis. To illustrate these aspects, Table 3 below shows percent differences in FGT measures between MI and SI welfare deflation at national and subnational level using the values of international poverty line for lower middle-income countries ($3.65, 2017 PPP) and upper middle-income countries ($6.85, 2017 PPP).

Table 3. Percent changes in poverty measures (MI-SI)

	$3.65				$6.85		
	H	**PG**	**PG2**		**H**	**PG**	**PG2**
National	-7.02	-7.81	-11.87		-0.29	-1.21	-2.67
Urban	-15.24	-17.37	-23.60		-1.90	-5.04	-8.32
Rural	22.18	24.95	23.42		4.41	10.47	15.12
Tehran metro Urban	-32.93	-36.50	-47.41		-5.72	-13.12	-20.04
Caspian Rural	2.93	2.34	2.61		0.84	1.17	1.71
Caspian Urban	-5.51	-4.33	-3.47		-5.07	-3.04	-3.67
Northwest Rural	8.46	15.88	0.00		12.86	13.37	12.55
Northwest Urban	0.85	0.23	-1.34		3.15	1.57	1.61
Northeast Rural	84.53	87.78	118.75		14.26	31.86	45.17
Northeast Urban	-2.39	-3.19	-3.07		-0.03	-1.43	-1.92
Central Rural	21.79	23.17	19.16		1.63	8.54	12.65
Central Urban	-7.92	-9.87	-11.84		-1.83	-2.67	-4.36
Southeast Rural	24.96	30.06	29.50		0.55	10.06	17.11
Southeast Urban	1.10	-0.36	-0.98		0.88	0.03	-0.03
Persian Gulf Rural	20.24	28.48	30.32		6.51	11.66	16.36
Persian Gulf Urban	-0.46	0.45	-0.85		0.30	0.41	0.82
Zagros Rural	54.26	21.21	0.00		11.50	18.78	23.91
Zagros Urban	2.62	2.61	3.90		0.35	1.41	1.76

Overall, differences in terms of poverty estimates between the two deflation approaches are sensitive to the value of the poverty line – this is what the analysis in Figure 2 suggests. Moreover, given that spatial price indexes are defined at the region/area level and that biases in urban/rural areas would go in opposite directions, differences in poverty estimates the national level tend to be relatively smaller compared to the ones at the subnational level. Considering the $3.65 poverty line, the MI approach would result in a 7 percent underestimation of poverty at the national level, which is the composite effect of a 15 percent underestimation of poverty in urban areas and a 22 percent overestimation of poverty in rural areas, compared to the SI approach. At the higher value of the poverty line ($6.85), the bias at the national level is negligible, while poverty in urban (rural) areas still results underestimated (overestimated) when using the MI approach compared to the SI approach.

In urban Tehran, compared to single index deflation, the MI approach would result in a 33 percent (5.7 percent) underestimation of poverty at the $3.65 ($6.85) poverty line. In rural Southeast the MI approach would result in a 25 percent (0.55 percent) overestimation of poverty at the $3.65 ($6.85) poverty line. Rural regions of Zagros and Northeast show an even higher poverty over-estimation bias when using the MI approach, reflecting a relatively larger differential between I_i^f and I_i^{nf} as well as relatively lower food budget shares.

The effect of different deflation methods on inequality measures is shown in Table 3. As is apparent in Figure 3, MI generally "mitigates" inequality in urban areas and "exacerbates" it in rural areas, with respect to SI, though magnitudes are very much dependent on the chosen measure. Considering these effects is relevant not only when inequality is the statistic of interest, but also when poverty dynamics is of interest (Ferreira 2012), particularly in settings where the mechanics underlying poverty changes are investigated (e.g., Datt and Ravallion 1992).

Table 3. Percent changes in inequality measures (MI-SI)

	Gini	**MLD**	**Theil**	**Atkinson** $\varepsilon = 0.5$	**Atkinson** $\varepsilon = 2$
National	-0.95	-2.25	-2.67	-2.25	-2.02
Rural	2.97	6.42	5.10	5.58	5.47
Urban	-2.10	-4.80	-4.85	-4.50	-4.35
Tehran metro Urban	-5.86	-12.87	-11.95	-11.58	-11.03
Caspian Rural	-0.05	-0.03	-0.38	-0.19	0.21
Caspian Urban	-0.81	-1.69	-1.63	-1.60	-1.40
Northwest Rural	0.61	0.62	-2.76	-0.64	1.32
Northwest Urban	0.49	0.98	1.09	0.98	0.66
Northeast Rural	5.17	11.66	10.99	10.77	10.57
Northeast Urban	-0.36	-0.75	-0.68	-0.68	-0.66
Central Rural	1.89	4.29	3.40	3.72	3.87
Central Urban	-0.80	-1.84	-1.65	-1.67	-1.74
Southeast Rural	9.68	19.62	19.04	18.62	13.31
Southeast Urban	-0.03	-0.11	-0.05	-0.07	-0.15

Persian Gulf Rural	6.26	12.94	14.32	13.00	10.07
Persian Gulf Urban	0.19	0.41	0.44	0.40	0.33
Zagros Rural	4.71	9.60	9.94	9.39	7.64
Zagros Urban	0.36	0.72	0.67	0.67	0.61

Differences between SI- and MI-based results presented so far are compatible with compensations in the re-ranking of individuals across the welfare distribution. The rearrangement of households along the distribution of expenditure can have important policy consequences. For this reason, in the final part of this section we focus on the non-anonymous distributional implications of switching from SI to MI, borrowing a few tools from the mobility literature.[14]

[14] We follow, by and large, the analysis presented by Ceriani, Olivieri and Ranzani (2022) in their evaluation of the distributional impact of imputed rent estimation.

Table 4. Probabilities of transition in and out of poverty when switching from SI to MI ($6.85 PL)

Area	% of poor individuals using SI who become non-poor using MI	% of non-poor individuals using SI who become poor using MI	number of individuals changing poverty status
National	3.7	1.2	1,620,300
Tehran metro Urban	9.3	1.7	686,288
Caspian Rural	0.2	0.5	12,923
Caspian Urban	5.4	0.1	62,248
Northwest Rural	6.2	3.8	133,883
Northwest Urban	0.0	0.9	49,406
Northeast Rural	5.3	3.6	102,378
Northeast Urban	0.7	0.2	22,455
Central Rural	2.0	1.1	49,644
Central Urban	2.2	0.1	62,538
Southeast Rural	4.8	8.8	197,767
Southeast Urban	0.5	1.1	40,399
Persian Gulf Rural	3.4	3.6	90,712
Persian Gulf Urban	0.3	0.2	14,949
Zagros Rural	4.2	2.9	95,337
Zagros Urban	0.6	0.2	18,733

First, we focus on the binary outcome of whether a household is labelled as poor or not, under the two alternative deflation scenarios. Taking the SI approach as a starting point, Table 4 shows the proportion of individuals that change their poverty status when the MI approach is adopted. Nationally, 3.7% of individuals who were poor under SI are no longer poor when using MI, while 1.2% of those who were non-poor under SI become poor when using MI. Overall, more than 1.5 million individuals in the country see their poverty status changed (from poor to non-poor or vice-versa) purely because of the choice between SI and MI. The single most relevant contributor to the size of these flows is Tehran, which is more populous to begin with, but several rural areas also display significant transitions in poverty status.

Table 5 gives a more detailed account of the re-ranking of households across the expenditure distribution: the population is sliced into decile groups, and a transition matrix is computed (the cells report the estimated probabilities of "jumping" from one expenditure group to another, again when switching from SI to MI). For example, in urban Teheran, a family that falls in the fifth decile group of real per capita expenditure when SI is used has a 57% probability of falling into the fifth decile again when using MI, a 25% chance of falling down to the fourth decile, and an 18% chance of jumping up to the sixth decile. In the rural Northeast in particular, jumps can span up to three groupings, showing that re-rankings can be considerable in magnitude.

The contents of Table 5 can be summarized using mobility indices. The Prais-Shorrocks index (Prais 1955, Shorrocks 1978), ranges between 0 (complete immobility) and 1.1 (complete origin independence) in this application, and is 0.33 for Teheran and 0.36 for the rural Northeast. The average jump index (Bartholomew 1973) ranges from 0 (complete immobility) to 1 (complete reversal), and it is 0.30 for Teheran and 0.36 for the rural Northeast, suggesting that average cross-decile movements due to the choice of MI vs. SI are substantive.

Table 5. Probabilities of transition across deciles of per capita expenditure, from SI (rows) to MI (columns)

		Teheran urban									
		Deciles MI									
		1	2	3	4	5	6	7	8	9	10
Deciles SI	1	0.86	0.13	0	0	0	0	0	0	0	0
	2	0.05	0.70	0.25	0	0	0	0	0	0	0
	3	0	0.08	0.63	0.29	0	0	0	0	0	0
	4	0	0	0.12	0.64	0.21	0	0	0	0	0
	5	0	0	0	0.25	0.57	0.18	0	0	0	0
	6	0	0	0	0	0.19	0.61	0.19	0	0	0
	7	0	0	0	0	0	0.29	0.59	0.12	0	0
	8	0	0	0	0	0	0	0.28	0.67	0.06	0
	9	0	0	0	0	0	0	0	0.14	0.84	0.02
	10	0	0	0	0	0	0	0	0	0.07	0.93
		Northeast rural									
		Deciles MI									
		1	2	3	4	5	6	7	8	9	10
Deciles SI	1	0.98	0.02	0	0	0	0	0	0	0	0
	2	0.25	0.73	0.02	0	0	0	0	0	0	0
	3	0	0.37	0.57	0.06	0	0	0	0	0	0
	4	0	0.02	0.27	0.57	0.14	0	0	0	0	0
	5	0	0	0.03	0.22	0.59	0.16	0	0	0	0
	6	0	0	0	0.01	0.22	0.54	0.22	0	0	0
	7	0	0	0	0	0.03	0.18	0.56	0.23	0	0
	8	0	0	0	0	0	0.04	0.13	0.6	0.23	0
	9	0	0	0	0	0	0	0.03	0.14	0.69	0.13
	10	0	0	0	0	0	0	0	0	0.12	0.87

6 Conclusions

The main aim of this paper is to clarify the implications of two alternative deflation strategies, which we call MI – multiple index – and SI – single index – on the welfare aggregate, and, based on those implications, to reach some conclusion of which strategy should be preferred.

Unless a household-level Paasche price index that has perfect coverage is available, real expenditures obtained using MI or SI will be different (Proposition 1). In these cases (arguably the majority of practical applications) the sign of the difference can be predicted under certain conditions, which are specified in terms of (i) the level of food and non-food prices, and (ii) the consumption pattern of households. In particular, for any given household, if the food price index is higher (lower) than the non-food price index, and its food budget share is higher (lower) than a certain threshold Δ, then the household's MI-adjusted real expenditure will be smaller than the SI-adjusted one (Proposition 2). Intuitively, with respect to SI, MI "penalizes" households that have a high budget share for the relatively expensive good. The budget share threshold Δ is a simple function of the values of the price indices: by computing it and comparing it with average budget shares in a sample, the analyst can predict the distributional effects of the SI vs. MI choice.

Given the systematic nature of the SI vs. MI differences, the paper asks whether it is possible to rank the two strategies, in terms of which one provides the better approximation of the welfare measure (be it Money-Metric Utility or Welfare Ratio). The analysis indicates that the two methods can indeed be ordered, albeit weakly. In fact, whenever a ranking can be established, SI proves to be the best strategy, somewhat surprisingly; in other words, SI weakly dominates MI (Propositions 3 and 4).

Through an empirical illustration, the paper shows that the SI vs. MI differences posited by the theory can have far-reaching implications in practice. We focus on spatial price differences in 2019 Iran, where the regional variation of non-food prices (here proxied by housing prices) is very pronounced, as compared to the variation of food prices. As a result, the non-food Paasche index is generally higher than the food Paasche in urban areas, and vice-versa in rural areas. Results indicate that when switching from SI to MI, national poverty decreases by about 7%, though this is the result of larger shifts in opposite directions happening at the local level: the price level and consumption pattern configuration in Iran implies a systematic overestimation of rural poverty (plus 22%) and a concurrent underestimation of urban poverty (minus 15%) when switching from SI to MI. In levels, the latter involves more people, given the density of urban areas; overall, more than 1.6 million people cross the poverty line in either direction, purely because of the choice between SI and MI.

It is worth asking whether this finding can be generalized to countries other than Iran. The answer depends on whether the price conditions observed in Iran can be thought of as a manifestation of some empirical regularity: is it true that, in general, non-food prices are lower in rural than in urban areas, while food prices are relatively homogeneous (as compared to non-food prices)? Such a pattern closely resembles one that is more commonly observed in international context, the Penn effect (the tendency for the cost of living to be higher in high-income countries), and its traditional explanation, due to Balassa (1964) and Samuelson (1964), which posits that spatial variation in the price of non-tradeable goods – in turn rooted in cross-country productivity differentials – is the main driver of these cost-of-living differences.[15]

A "within-country Penn effect", where the cost of living is higher in "rich" urban areas and lower in "poor" rural areas, mainly due to non-food prices (such as housing), does appear reasonable. Cheung and Fuji (2014) explicitly address this in a study of Japan, where they find that prices of non-tradeable goods are higher in more densely populated areas. They propose a Balassa-Samuelson-like mechanism to interpret the evidence: the agglomeration of economic activity induces productivity gains in urban areas through lower transport costs, externalities due to proximity, and specialization, which in turn boost wages (Ciccone and Hall, 1996); as a consequence, more workers are attracted to denser areas, and while this imparts pressure on the price of local non-tradeable goods, including housing and services, the same is not true for tradeable goods, such as food, whose prices tends to be equalized across regions.

Evidence on this phenomenon is sparser for low- and middle-income countries, given the well-known difficulties of observing non-food prices, especially in rural areas. Ravallion and van de Walle (1991) show that in Indonesia, both housing (hedonic) prices and food staple prices are higher in urban than in rural areas, although the variation is much larger for housing. Deaton and Dupriez (2011) observe that in India, the price level of food items is lower in rural areas, because of the large proportion of own-production and the effectiveness of food subsidies, while in Brazil, where these factors play a minor role, no clear urban-rural divide in food prices emerges: they suggest that food prices may vary less and less spatially as a country's income increases.

[15] The Balassa-Samuelson mechanism is based on the hypothesis that as economies develop, productivity gains tend to be located in sectors producing tradeable goods, while the non-tradeable goods sector remains relatively behind in terms of technological advances. This inflates the wage rate in the traded goods sector, and, assuming perfect mobility of labor across sectors, triggers an increase in the price level within the country. While the prices of traded goods tend to converge internationally, that is not case for non-traded goods prices, which are free to grow in the more productive/higher income economies.

Overall, though the literature provides too narrow an evidence base to make a definitive statement, the idea that many countries, especially in the low-and middle-income spectrum, experience spatial price patterns that are similar to those documented for Iran does not seem far-fetched. Housing prices play a major role in the overall non-food price level in poorer countries, and they are likely to spike in urban areas; food price differentials, on the other hand, may remain comparatively small. If so, then the implications of the analysis presented by this paper would be extremely relevant in terms of poverty profiling and targeting within a country, as well as in terms of the global poverty profile, with the potential for a widespread and systematic underestimation of urban poverty.

Appendix 1. Proofs

Proof of proposition 1

Observe that $(r, k, k_f, k_{nf}, \alpha, \alpha_f, \alpha_{nf}) = (H, K, K_f, K_{nf}, 0,0,0)$ implies:

$$I_h(H, K, 0) = \left[\sum_{j=1}^{K} w_j^h \left(\frac{p_j^0}{p_j^h}\right)\right]^{-1} = \frac{\sum_{j=1}^{K} q_j^h p_j^h}{\sum_{j=1}^{k} q_j^h p_j^0}$$

$$I_h^f(H, K_f, 0) = \left[\sum_{j=1}^{K_f} w_j^h \left(\frac{p_j^0}{p_j^h}\right)\right]^{-1} = \frac{\sum_{j=1}^{K_f} q_j^h p_j^h}{\sum_{j=1}^{K_f} q_j^h p_j^0}$$

$$I_h^{nf}(H, K_{nf}, 0) = \left[\sum_{j=1}^{K_{nf}} w_j^h \left(\frac{p_j^0}{p_j^h}\right)\right]^{-1} = \frac{\sum_{j=1}^{K_{nf}} q_j^h p_j^h}{\sum_{j=1}^{K_{nf}} q_j^h p_j^0}$$

It follows:

$$x_{MI}^h = \frac{x_f^h}{I_h^f(H, K_f, 0)} + \frac{x_{nf}^h}{I_h^{nf}(H, K_{nf}, 0)} = \frac{\sum_{j=1}^{K_f} q_j^h p_j^h}{I_h^f(H, K_f, 0)} + \frac{\sum_{j=1}^{K_{nf}} q_j^h p_j^h}{I_h^{nf}(H, K_{nf}, 0)} = \sum_{j=1}^{k} q_j^h p_j^0$$

However

$$\sum_{j=1}^{k} q_j^h p_j^0 = \frac{\sum_{j=1}^{K} q_j^h p_j^h}{I_h(H, K, 0)} = x_{SI}^h \Rightarrow x_{SI}^h = x_{MI}^h$$

■

Proof of proposition 2

Under the assumption $(r, k, k_f, k_{nf}, \alpha, \alpha_f, \alpha_{nf}) = (r, k, k_f, k_{nf}, \alpha, \alpha, \alpha)$, we can rewrite x_B^h and x_A^h as follows:

$$x_{SI}^h = \frac{w_h x_h}{I_i(h, k, \alpha)} + \frac{(1 - w_h) x_h}{I_i(h, k, \alpha)} = \frac{x_h}{I_i(h, k, \alpha)}$$

$$x_{MI}^h = \frac{w_h x_h}{I_i^f(h, k_f, \alpha)} + \frac{(1 - w_h) x_h}{I_i^{nf}(h, k_{nf}, \alpha)}$$

And

$$x_{SI}^h - x_{MI}^h = x_h \left[w_h \left(\frac{I_i^f - I_i}{I_i I_i^f}\right) + (1 - w_h) \left(\frac{I_i^{nf} - I_i}{I_i I_i^{nf}}\right)\right]$$

Assume $I_h^f > I_h^{nf}$, hence $x_{SI}^h - x_{MI}^h \neq 0$ and $x_{SI}^h - x_{MI}^h > 0$ if and only if the term in the squared bracket is positive which holds true if and only if:

$$w_h > \left(\frac{I_i - I_i^{nf}}{I_i^f - I_i^{nf}}\right)\frac{I_i^f}{I_i} = \Delta_i$$

It is immediate to verify that if $I_h^f < I_h^{nf}$ the term in the squared bracket is positive if and only if $w_h < \Delta_i$

■

Proof of the corollary

Under the assumption $\left(r, k, k_f, k_{nf}, \alpha, \alpha_f, \alpha_{nf}\right) = \left(r, k, k_f, k_{nf}, 1{,}1{,}1\right)$, all the indices are of the Laspeyres type and we can write $I_i = w_i I_i^f + (1 - w_i) I_i^{nf}$. Hence, the threshold value Δ_i simplifies to:

$$\Delta_i = w_i \frac{L_i^f}{L_i}$$

■

Proof of proposition 3

Under the assumption on preferences in the text, we know that $TCLI_{PK}^h > P_h(H, k)$ (Konüs, 1924) which, in turn, implies $MMU^h = x_h / TCLI_{PK}^h < x_h / P_h(H, k) = x_{SI}^h$. Hence $\left(x_{SI}^h - MMU^h\right) > 0$. Suppose now that $I_i^f > (<) I_i^{nf}$ and $w_h < (>) \Delta_i$, then $x_{MI}^h > x_{SI}^h$ and $\left(x_{MI}^h - MMU^h\right) > \left(x_{SI}^h - MMU^h\right) > 0$. Method (MI) tends to overestimate the real expenditures with respect to method (SI).

■

Proof of proposition 4

Under the assumption on preferences in the text, we know that $TCLI_{LK}^h < L_h(H, k)$, which implies $\widetilde{WR}^h = x_h / TCLI_{LK}^h > x_h / L_h(H, k) = x_{SI}^h$. Hence $\left(x_{SI}^h - \widetilde{WR}^h\right) < 0$. Suppose now that $I_i^f < (>) I_i^{nf}$ and $w_h > (<) \Delta_i$, then $x_{MI}^h < x_{SI}^h$ and $\left(x_{MI}^h - \widetilde{WR}^h\right) < \left(x_{SI}^h - \widetilde{WR}^h\right) < 0$. Method (MI) tends to underestimate the real expenditures with respect to method (SI).

■

Appendix 2. Sensitivity analysis

This appendix reports results obtained following an alternative approach for estimating the price variation of non-food items. In particular, the "single" Paasche index (SI approach) is constructed as in the equation below:

$$I_i = \left[\sum_{j=1}^{k_f} w_j^i \frac{p_j^0}{p_j^i} + w_r^i \frac{p_r^0}{p_r^h} + w_{nfnr}^i \right]$$

where j denotes each food item (belonging to the food commodity group), w_r^i denotes the budget share of housing expenditure, p_r denotes predicted rents, while w_{nfnr}^i denotes the budget share of *all* non-food/non-rent items, which are assumed not to vary spatially.

Correspondingly, the food and non-food price indices to be used in multiple index spatial deflation (MI approach) are estimated as follows:

$$I_i^f = \left[\sum_{j=1}^{k_f} w_j^i \frac{p_j^0}{p_j^i} \right]$$

$$I_i^{nf} = \left[w_r^i \frac{p_r^0}{p_r^h} + w_{nfnr}^i \right]$$

Figure A1. Food, non-food, and total price indexes

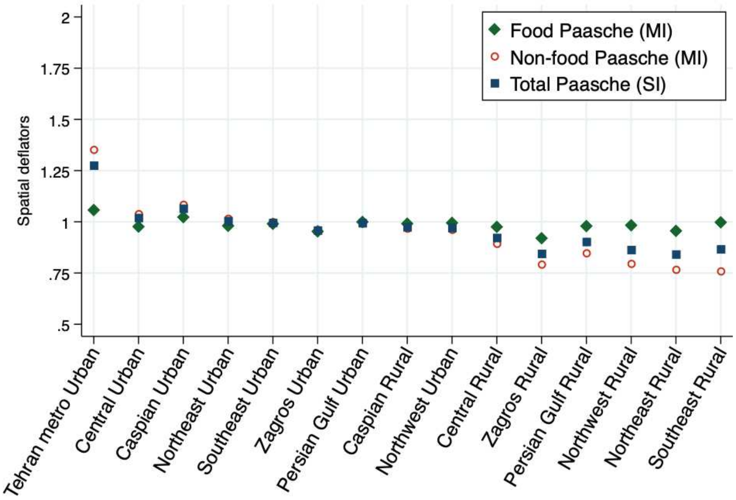
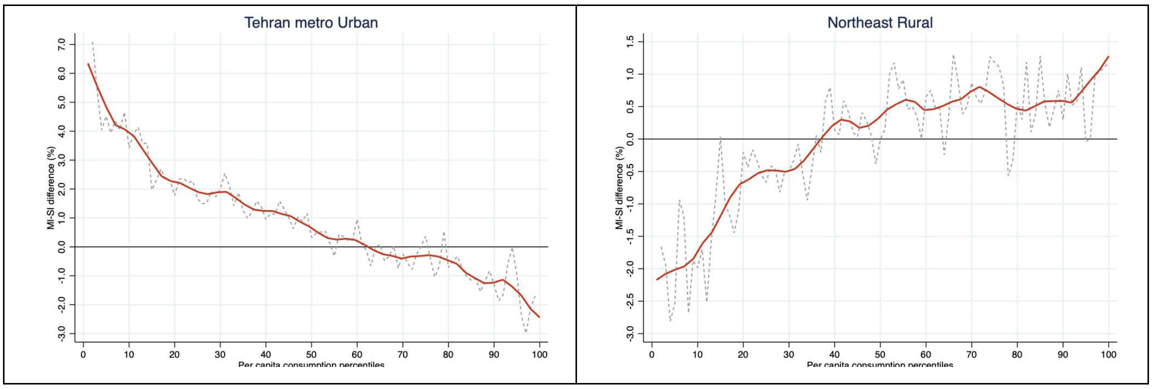


Figure A2. Percent difference in consumption aggregate deflated using MI and SI approaches, by percentile of the distribution in urban Tehran and rural Northeast



Note: The grey dashed line represents percentage differences at each percentile, while the solid red line is obtained via Kernel-weighted local polynomial smoothing.

Table A1. Percent changes in poverty measures (MI-SI)

	$3.65				$6.85		
	H	**PG**	**PG2**		**H**	**PG**	**PG2**
National	-0.26	0.00	0.00		-0.39	-0.78	-0.72
Urban	-4.40	-5.76	-7.34		-0.79	-2.18	-3.15
Rural	3.70	5.17	5.25		0.22	0.99	1.96
Tehran metro Urban	-19.51	-24.36	-32.74		-4.32	-9.52	-13.35
Caspian Rural	0.00	0.90	0.82		0.49	0.45	0.63
Caspian Urban	-0.58	-0.33	0.00		0.41	-0.39	-0.69
Northwest Rural	1.37	1.21	0.38		0.47	0.78	1.13
Northwest Urban	0.17	0.38	0.00		0.79	0.53	0.61
Northeast Rural	8.05	9.11	10.29		0.41	1.51	3.29
Northeast Urban	-0.60	-1.85	-2.78		0.44	-0.61	-0.90
Central Rural	5.25	6.41	5.73		0.52	1.76	2.79
Central Urban	0.00	-5.02	-5.93		-1.20	-1.32	-2.10
Southeast Rural	3.75	5.84	6.25		-0.79	0.62	2.03
Southeast Urban	-0.20	-0.06	-0.08		0.69	0.06	0.04
Persian Gulf Rural	4.06	5.06	5.14		0.81	1.58	2.21
Persian Gulf Urban	0.82	0.85	0.81		0.00	0.43	0.48
Zagros Rural	2.90	3.97	2.76		0.60	1.00	1.49
Zagros Urban	-2.26	-0.60	-1.02		-0.14	-0.31	-0.36

Table A2. Percent changes in inequality measures (MI-SI)

	Gini	MLD	Theil	Atkinson $\varepsilon = 0.5$	Atkinson $\varepsilon = 2$
National	-0.54	-1.03	-1.46	-1.14	-0.59
Rural	0.61	1.35	0.87	1.08	1.20
Urban	-0.78	-1.67	-1.88	-1.65	-1.31
Tehran metro Urban	-1.92	-4.38	-3.99	-3.88	-3.73
Caspian Rural	0.01	0.03	-0.04	-0.01	0.07
Caspian Urban	-0.23	-0.49	-0.49	-0.48	-0.40
Northwest Rural	-0.11	-0.43	-1.46	-0.78	-0.15
Northwest Urban	0.16	0.32	0.38	0.33	0.22
Northeast Rural	1.13	2.52	2.21	2.27	2.39
Northeast Urban	-0.22	-0.47	-0.48	-0.45	-0.41
Central Rural	0.48	1.08	0.83	0.94	0.98
Central Urban	-0.38	-0.85	-0.75	-0.77	-0.79
Southeast Rural	3.46	6.64	6.55	6.39	4.38
Southeast Urban	-0.04	-0.10	-0.07	-0.07	-0.09
Persian Gulf Rural	1.60	3.21	3.59	3.26	2.50
Persian Gulf Urban	0.04	0.08	0.08	0.08	0.07
Zagros Rural	0.85	1.67	1.71	1.62	1.30
Zagros Urban	-0.05	-0.11	-0.10	-0.10	-0.08

Table A3. Probabilities of transition in and out of poverty when switching from SI to MI ($6.85 PL)

Area	% of poor individuals using SI who become non-poor using MI	% of non-poor individuals using SI who become poor using MI	number of individuals changing poverty status
National	1.2	0.3	482,400
Tehran metro Urban	5.4	0.2	187,498
Caspian Rural	0.0	0.2	4,254
Caspian Urban	0.8	0.3	18,189
Northwest Rural	1.5	1.7	51,295
Northwest Urban	0.0	0.3	15,823
Northeast Rural	1.9	1.8	48,837
Northeast Urban	0.0	0.2	10,411
Central Rural	0.5	0.5	18,982
Central Urban	1.5	0.1	41,593
Southeast Rural	1.1	1.0	33,777
Southeast Urban	0.0	0.5	13,622
Persian Gulf Rural	0.8	1.4	28,674
Persian Gulf Urban	0.0	0.0	0
Zagros Rural	1.1	1.2	35,618
Zagros Urban	0.3	0.1	9,783

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