ELSEVIER

Journal of Development Economics
Vol. 68 (2002) 187–203

JOURNAL OF
Development
ECONOMICS

www.elsevier.com/locate/econbase

Prices and living standards: evidence for Rwanda

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Received 30 August 1999; accepted 30 July 2001

Abstract

We review the explanations of the statistical relationship between spatial price indices and real living standards. Then, using data from several seasons in rural Rwanda, we show that these variables are negatively associated, hinting at price discrimination against the poor. In that case, policies permanently improving of market functioning may simultaneously improve efficiency and equity.

Moreover, under a hypothesis of weak association of nominal living standards and price indices, we derive simplified formulae for social welfare indicators. These formulae depend only on a small number of sample statistics obtainable from separate publications for prices and living standards.

Résumé

La compréhension du lien statistique entre prix et niveaux de vie est essentielle á l'analyse de bien être. Nous passons d'abord en revue les explications possibles de cette relation. Nous montrons ensuite que lorsque les pauvres font face ä une discrimination en prix, le lien entre niveaux de vie nominaux et indices de prix est atténué. Sous cette hypothése, nous dérivons des formules simplifiées d'indicateurs populaires de bien-être social. Ces formules dépendent seulement d'un petit nombre de statistiques simples pouvant être obtenues á partir de publications séparées pour les prix et les niveaux de vie.

A partir de données du Rwanda nous montrons que les formules simplifiées fonctionnent bien. En revanche, indices de prix et niveaux de vie réels sont négativement associés, ce qui suggére la présence de discrimination en prix á l'encontre des pauvres. Par conséquent, améliorer de facon

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PII: S0304-3878(02)00011-1

permanente le fonctionnement spatial des marchés peut augmenter simultanément l'efficacité et l'équité.

JEL classification: I32 (measurement and analysis of poverty); O15 (income distribution); D31 (personal income distribution)
Keywords: Rwanda; Prices; Living standards

1. Introduction

Understanding the statistical link between prices and household living standards is central to social welfare analysis. Indeed, not only is the real living standard indicator (y) composed of two fundamental variables, the nominal living standard (x) and the price index (P), but also the geographical correlation between prices and living standards matters for social policy. As Sen (1981) and Stern (1989) stress, the correlation between food prices and income is crucial for understanding the capacity of the poor in LDCs to feed themselves. Moreover, even policies that strengthen incentives to expand food production through higher food prices may result in severe hardship for the poor at least in the short run (Pinstrup-Andersen, 1985).

Although prices of specific products are likely to be important for producer incomes, it is not known how the general level of prices in an area is linked to the real living standards of households in the same location. Using data for a few US cities, Black (1952) find price levels positively correlated with income levels. In Indonesia, Ravallion and van de Walle (1991), Bidani and Ravallion (1993), and Ravallion and Bidani (1994) find that cost-of-living indices based on market prices tend to be higher in urban areas and food prices to be lower in poorer provinces (therefore with strong correlations with nominal expenditures). What is less clear is if the same systematic link between x or y, and P occurs at household level and in the rural areas of LDCs, where most of the poor are concentrated, and which are the object of specific price and welfare policies. In three Indian villages, Rao (1997, 2000) finds that among 140 potter families poor households pay higher prices because of liquidity constraints forcing them to purchase goods in small quantities. He also found that family size is negatively correlated with prices, but that the amount of land owned shows a positive relationship.

The seasonality is also often implemented. Because prices and incomes in rural areas considerably fluctuate across seasons, we calculate quarterly indicators rather than annual indicators of prices and living standards. Then, we can analyse the link of household living standards and prices at different seasons instead of using annual means as it is generally done.

What is the statistical link of price indices and living standards in rural LDCs? What are the consequences of this link for welfare analysis? The aim of this article is to study these questions by investigating the geographical association of living standards and spatial price indices in rural Rwanda. We discuss the theoretical relationship between price indices and living standards in Section 2. We present the data in Section 3. We discuss the results of price index regressions in Section 4. In Section 5, we show how welfare indicators can be simplified because P and x are often likely to be weakly statistically linked. The simplified indicators are shown to perform well with our data. Finally, Section 6 concludes.

2. The relationship of price indices and living standards

Price indices and real living standards may be statistically linked on several grounds. Firstly, this link may arise from the methods used in defining the indicators. Using unit-values (ratios of consumption values by consumption quantities) instead of observed market prices may entail spurious correlations between consumption levels and unit-values[1] and therefore between living standards and price indices. However, market prices should be used when possible, and we shall not investigate these spurious correlations.

Secondly, because households have different preferences and different incomes, their individual true price indices must differ owing to variations in the chosen baskets of goods, even when they face the same prices. The impact of such substitution effects has already been studied and estimated true price indices can be derived from the estimation of a demand system (as in Braithwait, 1980; Ravallion and van de Walle, 1991; Slesnick, 1993). If the poor consume proportionally fewer expensive goods than the rich, then true price indices and real living standards are positively correlated.

Thirdly, local supply and demand effects may result from the imperfect spatial integration of markets. Rich households may live in a different economic environment from poor households and have different structures of consumption and production. In this situation, local prices in poor areas would be different from those in rich areas. The spatial link of P and y resulting from a general equilibrium is generally of ambiguous direction. However, in predominantly rural economies, the rich tend to be net suppliers of basic agricultural products and the poor to be net demanders (Saith, 1982; Ravallion, 1987). Then, areas where the price of the staple food is higher may be associated with a worse income distribution. Because the agricultural products consumed by the poor constitute a major component of the price index, a negative correlation between y and P may arise. Fourthly, migrations or household locations may relate y and P, if for example poor households move to areas where prices are lower.

Fifthly, various market imperfections influence the relation between y and P. Transaction costs are relatively less important for the rich than for the poor. The poor may face higher prices because of liquidity constraints forcing them to buy goods in too small a quantity (Musgrove and Galindo, 1988; Rao, 1997, 2000) or at suboptimal periods. The poor may also bear higher searching costs for lack of modern means of transport, or because they live far from commercial centres (Greenhult et al., 1987). Since the volume and the bought share of their consumption are lower than that of the rich, they may gain less from searching for the lowest possible price (Alcaly, 1976; Anglin and Baye, 1987).

In regions where the poor live, they may not have access to a local credit market. In this case, the local price level may be higher, reflecting the inefficiency coming from insufficient

[1] Deaton (1988, 1990) has developed a method for removing these spurious correlations arising from quality differences.

intertemporal adjustments. Areas with higher transaction costs may be characterised by higher price levels, because of mark-up mechanisms similar to the inflation by the costs, and by less efficient economic interactions associated with lower incomes.

Also, producers in high prices areas may have relatively higher production levels, responding to their misled perception of advantageous output prices. This would cause a relatively higher local labour demand, therefore relatively higher wage rates, themselves causing higher price levels through mark-up mechanisms, and an increase of labour supply. In this hypothesis, producers perceive the higher nominal wage as a lower real wage, while workers perceive higher real wages. There, y and P would be positively associated before adjustment of perceptions to actual prices.

Finally, the poor may face deliberate price discrimination. Local monopolists (Holahan, 1975) or fragmented duopolists (Basu and Bell, 1991) can design price discriminating schemes based on demand elasticities varying with household income (or distance to selling sites that would be correlated with income). However, general models of price discrimination imply ambiguous consequences as to the impact on prices (Nahata et al., 1990). This prevents us from deriving safe conclusions as to the link of y and P in that case.

The numerous competing explanations for the link of y and P, and the fact that the data cannot identify most of them, lead us to focus on the statistical relationship of y and P, which we estimate in Section 4. Next, we present the data for this estimation.

3. The data

We use 1983 data from Rwanda. Our aim is not so much to say something specific about Rwanda, but to avail of a data set that is appropriate to address the issue concerned. In this regard, the data have several crucial characteristics: the measurement errors of living standard indicators are likely to be very small; there exist local and seasonal prices; and the information is available for several quarters, which allows the investigation of spatial correlation of price indices and living standards without mixing them with seasonal variations that are important for these variables. Rwanda is a very poor country, with per capita GNP of 1983 US $270 per annum. More than 95% of the population of 5.7 million in 1983, lives in rural areas.[2] Agriculture is the mainstay of the economy.

We use data taken from the Rwandan national budget-consumption survey, conducted in the rural part of the country from November 1982 to December 1983.[3] The agricultural year 1982–1983 was a fairly normal year in terms of climatic fluctuations. The collection of the consumption data was organised in four rounds, corresponding to four quarters (A, B, C, D)[4] of the agricultural year 1982–1983. Two hundred seventy households were surveyed about their budget. Owing to missing values, only 256 households are used in the estimation.

[2] Bureau National du Recensement (1984).

[3] See Ministere du Plan (1986). The main part of the survey was designed with the help of INSEE (French national statistical institute). The author was himself involved in this project as a technical adviser from the French Ministry of Cooperation and Development.

[4] Round A: 01/11/1982 until 16/01/1983. Round B: 29/01/1983 until 01/05/1983. Round C: 08/05/1983 until 07/08/1983. Round D: 14/08/1983 until 13/11/1983.

The consumption indicators are of exceptional quality relative to what is obtained in most household surveys. This is first due to a very intensive collection. The volume of all containers present in the household was measured in every household in addition to the traditional measurement units, so as to obtain accurate information about quantities of goods. Every household was visited at least once a day during 2 weeks for every quarter. Daily and retrospective interviews and recordings of food weights were carried out. Every household had to fill a diary between the quarterly survey rounds. The set of surveyed topics was very broad, and several questions have been repeatedly treated by different collection methods. This enabled better collection, a thorough cleaning of the data and sophisticated verification algorithms based on these many redundancies present in the data. Finally, we have calculated seasonal consumption indicators on algorithms aimed at reducing the measurement errors from the optimal combinations of questionnaires.

Geographical and seasonal price dispersions are considerable in Rwanda.[5] We dispose of elementary price indicators of the main categories of goods, for every season and every cluster of the sample. The prices of each category of goods are represented by the price of the main product. Then, comparing prices across seasons and clusters can be done without quality bias since the same product is used. Another element of the calculation of Laspeyres price indices is the estimation of a balanced composition of the aggregate consumption for rural Rwanda by categories of goods (Muller, 1992).

Theoretical price indices are ratios of cost functions representing the household preferences, which have been utilised in applied welfare studies.[6] However, in practice, they are generally approximated by Laspeyres or Paasche price indices. In this paper, the price correction is carried out using a Laspeyres price index (P_{it}) specific to each household and each quarter, in which the basis is the annual national average consumption weighted by the sampling scheme. We chose this index because it is the most common in applied work. It has also the advantage of being insensitive to estimation noise from a demand system and being unbiased by symmetric measurement noise on prices as opposed to the Paasche index. It is a first-order approximation to Blackorby and Donaldson's (1987) welfare ratio (expenditure divided at a point on the cost function at reference utility). Note that it eliminates most substitution effects from the link of y and P. To control for this and for the robustness to alternative price indices, we conduct the same analyses with a household-specific Paasche index (own consumption bundle at common price vector). Expenditure deflated by this index is a first-order approximation to money-metric utility (although money-metric utility is a somewhat fragile reference in such an imperfect market situation). All our qualitative results are confirmed with this Paasche price index (and other variants).

The real living standard indicator for household i at period t is $y_{it} = c_{it}/(\mathrm{es}_i P_{it}) = x_{it}/P_{it}$, where c_{it} is the value of the consumption of household i at quarter t, x_{it} is the nominal

[5] See O.S.C.E. (1987) and Muller (1988).

[6] Muellbauer (1974), Glewwe (1990), Ravallion and van de Walle (1991), and Grotaert and Kanbur (1996). Although the error in measuring the money metric utility by not using a true price index is a topic of interest that has already been well studied, it is not the focus of our investigation. One difficulty with using predicted true price indices derived from demand system estimates is that the prediction error attached to these price indices can be considerable.

Table 1
Means and standard deviations of the main variables

	RC	RPCC	NPCC	P	H
Annual	51176 (24985)	10613 (5428)	10905 (5355)	1.048 (0.063)	
A	13521 (9527)	2750 (1701)	2995 (1826)	1.108 (0.129)	1.093
B	13232 (8192)	2702 (1620)	2539 (1475)	0.953 (0.101)	0.941
C	13452 (8249)	2850 (1968)	2902 (1834)	1.047 (0.131)	1.029
D	10969 (6092)	2310 (1511)	2468 (1524)	1.084 (0.097)	1.075

RC = real consumption per household (in Frw); RPCC = real per capita consumption (in Frw); NPCC = nominal per capita consumption (in Frw); P = Laspeyres price index; H = harmonic mean of the price index.
Standard deviations are in parentheses.
In 1983, the average exchange rate was 100.17 Frw for one 1983 US \$, i.e. 60.16 Frw for one 1999 US \$ (sources: IMF, Penn Tables).

standard of living of household i at quarter t, es_i is the adult-equivalent scale of household i and P_{it} is the price index associated with household i and quarter t. Four equivalence scales have been used, which are shown in Muller (1999), where various association statistics deliver results consistent with the correlation statistics of this paper. Because the qualitative results for these different scales are similar, we focus on the case of the per capita consumption. We proceed with the empirical study of the statistical link of y and x.

Table 2
Mean price indices

Mean price indices by quintile

Type of y for the quintile definition	Quarter	$Q1$	$Q2$	$Q3$	$Q4$	$Q5$
Quarterly	A	1.165	1.131	1.093	1.090	1.063
Quarterly	B	0.986	0.978	0.967	0.931	0.908
Quarterly	C	1.091	1.058	1.053	1.052	0.982
Quarterly	D	1.105	1.092	1.110	1.081	1.032
Annual	A	1.134	1.085	1.115	1.094	1.113
Annual	B	0.963	0.979	0.953	0.952	0.918
Annual	C	1.069	1.064	1.060	1.048	0.995
Annual	D	1.099	1.089	1.092	1.071	1.070

$Q1$–$Q5$ are the five quintiles

Mean price indices for the chronically rich the chronically poor

Quarter	CR0	CR1	CP0	CP1
A	1.106	1.121	1.104	1.145
B	0.959	0.917	0.953	0.952
C	1.054	1.007	1.044	1.076
D	1.084	1.086	1.085	1.079

CR0 = households that are not chronically rich; CR1 = chronically rich households; CP0 = households that are not chronically poor; CP1 = chronically poor households.

4. The link of real living standards and price indices

Table 1 presents the mean and standard deviation, for each quarter and for the year, of the real consumption, the real and the nominal per capita consumption and the Laspeyres price index. The independence of real living standards and price indices is rejected at 5% level by the χ^2 tests. Tables 2 and 3 show mean price indices and regressions of P against y, for various living standard classes. Fig. 1 shows the nonparametric regressions

Table 3
Price index regressions

Coefficients of the real living standards in the regression of the price index

Type	Quarter	OLS	Robust
Level	A	– 0.187E – 4	– 0.175E – 4
Level	B	– 0.144E – 4	– 0.144E – 4
Level	C	– 0.217E – 4	– 0.235E – 4
Level	D	– 0.164E – 4	– 0.186E – 4
Log	A	– 0.0611	– 0.0565
Log	B	– 0.0522	– 0.0485
Log	C	– 0.0650	– 0.0618
Log	D	– 0.0264	– 0.0304

These regressions are estimated for level and double-log specifications. The shown coefficients are all significant at 5% level.

Regressions in level by quintile (coefficient to multiply by 10^{-4})

Type	Quarter	*Q*1	*Q*2	*Q*3	*Q*4	*Q*5
Quarterly	A	– 1.09 *	– 1.45	– 1.18	– 1.20**	0.00158
Quarterly	B	– 0.33	– 2.33 **	– 1.40	– 0.39	0.190
Quarterly	C	– 0.038	– 0.28	– 0.197	0.40	0.130
Quarterly	D	0.56	– 0.64	2.26	0.88	0.285
Annual	A	– 0.596 *	– 0.768 *	– 0.342**	– 0.295 *	– 0.225 *
Annual	B	– 0.165	– 0.404	– 0.421 *	– 0.358 *	– 0.040
Annual	C	– 0.192	– 0.927 *	– 1.38	– 0.166	– 0.244 *
Annual	D	– 0.359	– 0.070	– 0.536 *	– 0.423 *	– 0.140 *

Regressions in logarithm by quintile

Type	Quarter	*Q*1	*Q*2	*Q*3	*Q*4	*Q*5
Quarterly	A	– 0.0796 *	– 0.220	– 0.277**	– 0.357**	– 0.00789
Quarterly	B	– 0.0235	– 0.397 *	– 0.315	– 0.0959	0.0826
Quarterly	C	0.00616	– 0.0252	– 0.050	0.120	0.0714
Quarterly	D	0.00403	– 0.0629	0.405	0.220	0.1211
Annual	A	– 0.0610 **	– 0.169 *	– 0.917 *	– 0.0809 *	– 0.099 *
Annual	B	– 0.0227	– 0.0817	– 0.102 *	– 0.101 *	– 0.0182
Annual	C	– 0.0228	– 0.154 *	– 0.0564	– 0.0392	– 0.141 *
Annual	D	– 0.0218	0.00238	– 0.0931 *	– 0.0310	– 0.0263

* Significant at 5% level.

** Significant at 10% level.

(i.e. curves) of P against y at each quarter. We comment on the estimation results separately for the whole population, for the quintiles of quarterly y, for quintiles of annual y, and finally by separating chronically rich and chronically poor households. All these results are qualitatively confirmed by using Paasche indices instead of Laspeyres indices, showing that they are unlikely to be related to substitution effects.

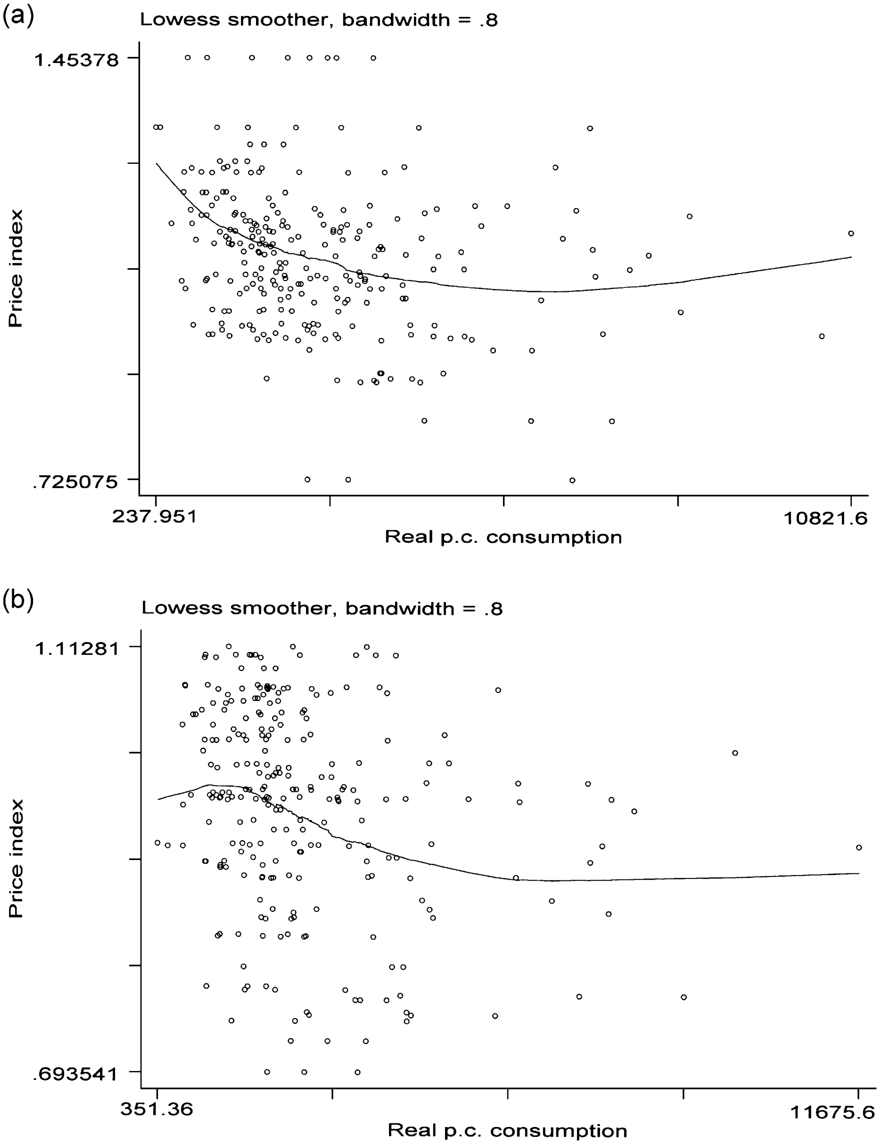
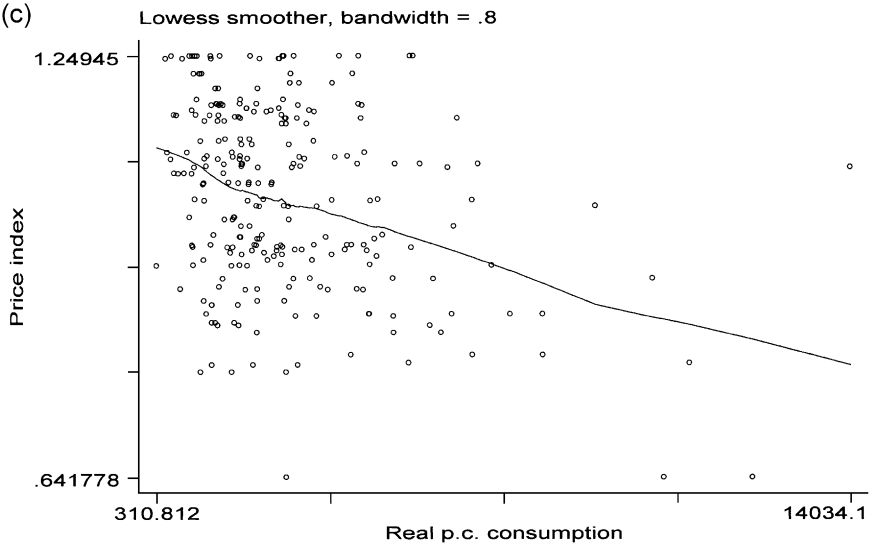


Fig. 1. (a) Regression curve at quarter A. (b) Regression curve at quarter B. (c) Regression curve at quarter C. (d) Regression curve at quarter D.

(c)



(d)

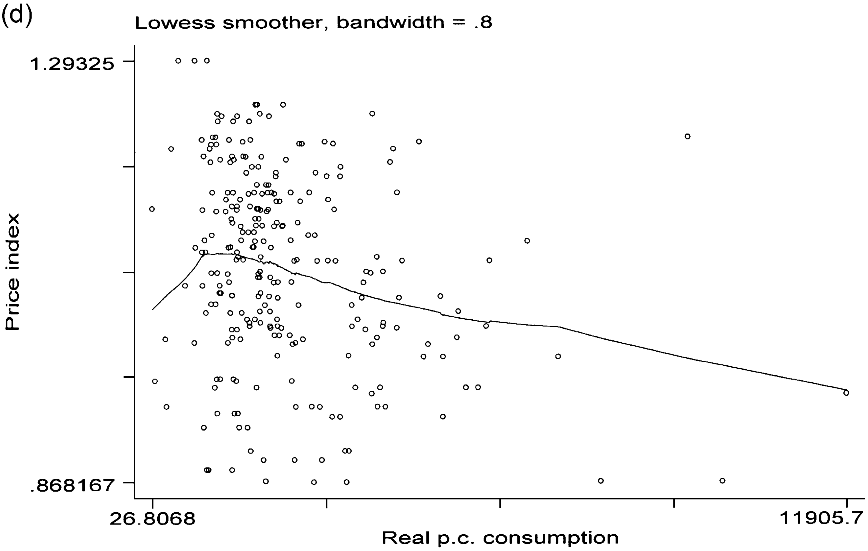


Fig. 1 (*continued*).

(a) *For the whole population*: The regression curves show that at all quarters price level decreases when per capita consumption rises.[7] Very significant negative links between P and y are found both for regressions in level and in logarithms. This is confirmed by robust

[7] The extremities of curves may not be significant because their estimates are based on a small sample of observations. For example, the anomaly at quarter D for extremely poor households is a statistical artefact.

regressions using an iterative combination of Huber and Biweights estimators (Huber, 1964; Beaton and Tukey, 1974) controlling for some measurement errors.

(b) *By quintiles of quarterly living standards*: For all quarters, the mean price index decreases when the quintile of quarterly y rises, except for the third quintile at quarter D. For quarterly quintiles, the regression coefficients of y are much less frequently significant at 5% level than for the whole population, although when they are, their sign is negative.

(c) *By quintiles of annual living standards*: The maximal mean price index at each quarter corresponds to the first or the second quintile of annual y. The regression coefficients of y are negative for all quintiles and quarters, and much more often significant at 5% level than with the quarterly quintiles.

(d) *Separating the chronically rich or the chronically poor*: Only 27 households out of 265 remain during all quarters in the fourth or fifth quintile of quarterly y. These "chronically rich households" face lower prices in quarters B and C. Similarly, only 27 households out of 265 remain at all quarters in the first or second quintile of quarterly y. In quarters A and C, these "chronically poor households" face higher prices, but lower ones in quarter D.

(e) *Remarks*: The less significant link when using the quintiles of quarterly y rather than the quintiles of annual y suggests several observations. First, the differences in prices affecting households identified by annual y may be associated to their "permanent living standard". For example, poor households may live far from the main market and transaction sites because migration to better location is costly, or/and they may be permanently vulnerable to price discrimination by sellers. Second, seasonal phenomena are unlikely to explain the link, otherwise the link would be more significant for quintiles of quarterly y. The mean price indices and the regression curves show that the negative link of P and y is continuous across the distribution of y rather than a clear-cut discrimination against the chronically poor or in favour of the chronically rich.

Endogeneity of observed quarterly consumptions and observed quarterly prices cannot explain the link because if endogeneity were important, the coefficients calculated with quarterly quintiles would be larger than with annual quintiles. We also control for endogeneity by using two-stage least square estimates (where instrumental variables for living standards are the household composition and the land) that deliver similar results. Finally, other household characteristics, such as land, distance to market and socio-demographic variables (age, gender, ethnic group and education of the head, household size) have been added to the regressions, but are almost always insignificant at 5% level. This largely rules out problems related to accounting for demographic characteristics in the definition of living standards.

Substitution effects cannot explain the observed associations since they are largely ignored by Laspeyres price indices. Moreover, using Paasche indices instead does not eliminate the negative link. Migrations could be an explanation, although one usually expects the rich to dwell in places where prices are higher. Because of strong market imperfections hinted at by the large levels of own-consumption and transfers in kind, there may be local supply/demand and local general equilibrium effects, although the direction of these effects could not be identified. Because poor households often tend to be net demanders of agricultural products that are the main components of the price index and rich households tend to be net suppliers, y and P tend to be negatively correlated. However, we did not find significant correlation between household agricultural surplus and P. These results do not support the previous conjecture based on net supply effects.

Many unobserved market imperfections constitute alternative explanations. Indeed, the presence of transaction costs, liquidity constraints and limited access to credit, suboptimal spatial and intertemporal price adjustments may all explain the negative link of y and P. Another explanation would be deliberate price discrimination by sellers, although nothing restricts this discrimination from being unfavourable to the poor. In contrast, imperfect perceptions by agents of spatial price differences would rather yield a positive correlation of y and P. Note that the continuity of the link of y and P across the distribution of y makes explanations based on transaction costs or liquidity constraints less convincing because of fixed costs or thresholds beyond which the link should disappear. A word of conclusion on this empirical section: The spatial negative link of y and P is well established and robust, although it cannot be well explained by any of the theories that we reviewed.

5. Simplifications in welfare analysis

We now show that x and P are likely to be more loosely related than y and P, notably when the latter ones are negatively linked. We develop this case firstly since it corresponds to our data. Let us compare: (i) $\ln P = \alpha_1 + \beta_1 \ln y + u_1$ and (ii) $\ln P = \alpha_2 + \beta_2 \ln x + u_2$, where α_1, α_1, α_2, β_1 and β_2 are parameters and u_1 and u_2 are error terms. The OLS estimators of coefficients β_1 and β_2 (respectively $\hat{\beta}_1$ and β_2) correspond to our correlations of interest, respectively between $\ln P$ and $\ln y$, and between $\ln P$ and $\ln x$. Because of the similarity of these variables, one expects that the two coefficients have the same sign, which is confirmed in our data. We obtain: $\hat{\beta}_1$=(Cov(ln y, ln P))/(V(ln y)) and β_2=(Cov(ln x, ln P))/(V(ln x)), where Cov($\cdot$) and V($\cdot$) are the empirical covariance and variance operators. Then, $\hat{\beta}_2 = \gamma \hat{\beta}_1$, where

$$\begin{aligned}
\gamma &\equiv \frac{\mathrm{Cov}(\ln x, \ln P)}{V(\ln x)} \frac{V(\ln y)}{\mathrm{Cov}(\ln y, \ln P)} \\
&= \frac{\mathrm{Cov}(\ln y + \ln P, \ln P)}{V(\ln y + \ln P)} \frac{V(\ln y)}{\mathrm{Cov}(\ln y, \ln P)} \\
&= \frac{V(\ln y)}{V(\ln y + \ln P)} + \frac{V(\ln y) \cdot V(\ln P)}{V(\ln y + \ln P) \cdot \mathrm{Cov}(\ln y, \ln P)} \\
&\le \frac{V(\ln y)}{V(\ln y + \ln P)} \text{[because } \mathrm{Cov}(\ln y, \ln P) < 0] \\
&\le 1 \text{ [by stochastic dominance]}
\end{aligned}$$

The latter step of stochastic dominance is valid because typical price and income data correspond to small correlations of price indices and incomes. It is also possible to study directly the formula of γ under reasonable hypotheses on the distribution of $\ln y$ and $\ln P$. We carried out simulations showing that $\gamma < 1$ for realistic correlations and variances of $\ln P$ and $\ln y$. Therefore, there may exist attenuation of the link of x and P. Naturally, the attenuation does not mean that the link between x and P is necessarily insignificant. We now investigate the consequences on social welfare measures of a weak link.

Beyond the above attenuation, a weak link of x and P may also occur because of strong market imperfections disconnecting incomes and prices. For example, in the administrated economy of Russia during the latter part of the 1980s, the evolutions of nominal wages and price levels have been found to be unrelated (Koen and Phillips, 1993). We now exploit the weak statistical link of x and P to produce simplifications in social welfare analysis. Under independence, the marginal distributions are sufficient parameters (of infinite dimensions) of the joint distribution. In that case, any functional defined in terms of the joint distribution is a functional of the two marginal distributions only. However, this is useful only if explicit formulae can be exhibited, which we do for welfare indicators. Moreover, we relax the assumption of independence.

All that is required for our simplifications is:

$$\int\int k(x,P)\mathrm{d}F(x,P) = \int\int k(x,P)\mathrm{d}F_1(x)\mathrm{d}F_2(P) \tag{5.1}$$

where k is any kernel function associated with any integral component involved in the formula of the welfare indicator of interest, F is the joint c.d.f. of x and P, and F_1 and F_2 are the marginal c.d.f. of x and P. We call this condition 'weak statistical association of x and P for the welfare measure', in short 'weak association of x and P'. A sufficient condition is that x and P are independent. The weak association of x and P could be tested for any specific welfare measure by using joint data on prices and living standards. This condition for a given welfare indicator can be considerably weaker than the independence, but is useful only for dealing with this indicator. Even when the link of x and P is statistically significant, we do not expect it to be high. Then, the case of weak association provides useful insight as an approximation.

Let us consider an integral component of a welfare indicator defined by its kernel function $k(x, P) = f_1(x)f_2(P) + f_3(x) + f_4(P) + K$, where f_1, f_2, f_3, f_4 are integrable functions and K is a constant. This functional form is frequent in welfare analysis. Then, we have under independence $W = Ef_1(x)\cdot Ef_2(P) + Ef_3(x) + Ef_4(P) + K$, where E denotes the expectation over the relevant populations. This result covers most cases of interest.

Let us consider a few examples. When x and P are weakly associated (for the mean here, and below unmentioned for each relevant welfare index), $\bar{y} = \int_0^{+\infty}\int_0^{+\infty} \frac{x}{P}\mathrm{d}F_1(x)\mathrm{d}F_2(P) = \frac{\bar{x}}{H}$, where $\bar{x}$ is the arithmetic mean of living standards and H is the harmonic mean of price indices. The variance of y, ${\sigma_y}^2$, and the variance of the logarithms, ${\sigma_{\ln y}}^2$, are simple inequality measures. When x and P are weakly associated, we obtain:

$$\sigma_y^2 = E(x^2)\cdot E\left(\frac{1}{P^2}\right) - \left(\frac{\bar{x}}{H}\right)^2 \text{ and } \sigma_{\ln y}^2 = \sigma_{\ln x}^2 - \sigma_{\ln P}^2 \tag{5.2}$$

The case of the coefficient of variation, CV, is a direct corollary of the cases of the mean and the variance. The Theil index is defined as $I_T \equiv \int\int (x/P)/(\mu(F)) \ln[(x/P)/(\mu(F))] \mathrm{d}F(x,P)$, where $\mu(F) \equiv \int\int x/P \mathrm{d}F(x,P)$. Under weak association of x and P, we have:

$$I_T = \frac{E(x\ln(x))}{\bar{x}} - \ln(\bar{x}) - H \cdot E(\frac{\ln P}{P}) + \ln(H) \tag{5.3}$$

The Atkinson index with parameter ε is defined as $I_A^{\varepsilon} \equiv 1 - (1/(\mu(F))[\int\int (x/(P))^{1-\varepsilon} \mathrm{d}F(x,P)]^{1/(1-\varepsilon)}$, which gives under weak association of x and P:

$$I_A^{\varepsilon} = 1 - \frac{H}{\bar{x}} \left[E(x^{1-\varepsilon}) \cdot E\left(\frac{1}{P^{1-\varepsilon}}\right)\right]^{1/(1-\varepsilon)} \tag{5.4}$$

The Generalised Entropy inequality measure with parameter α is defined as $I_{\mathrm{GE}}^{\alpha} \equiv (1/(\alpha^2 - \alpha))\int\int ((x/P)/\mu(F))^{\alpha} - 1 \mathrm{d}F(x,P)$, which gives under weak association of x and P:

$$I_{\mathrm{GE}}^{\alpha} \equiv \frac{1}{\alpha^2 - \alpha} \left[E(x^{\alpha}) \cdot \left(\frac{H}{\bar{x}}\right)^{\alpha} \cdot E\left(\frac{1}{P^{\alpha}}\right) - 1\right] \tag{5.5}$$

When the building blocks of poverty or inequality measures do not straightforwardly simplify under weak association, one can use Taylor expansions to approximate the kernel functions and to apply the same method. In all the considered cases, only a small number of price statistics are necessary to summarise the effect of the price distribution for the welfare indicator. The simplified versions of welfare indicators are more complicated than what is done in practice by wholly omitting price effects. However, even when the basic price statistics are not directly available, e.g. the harmonic mean of price indices, they are very easy to calculate. Their publication could be routinely undertaken by statistical offices at very low cost. They can also be extrapolated from the available information by using a distribution model for P.

Let us now confront these formulae with our data. The χ^2 tests cannot reject the independence of price indices and nominal living standards at 5% level. In this case, x and P are weakly associated for all welfare indicators and we can use the simplified formulae of welfare statistics. The considered statistics are the mean, the standard deviation, the coefficient of variation, Theil inequality index, a Generalised Inertia index (with $\alpha = 2$) and an Atkinson index (with $\varepsilon = 0.5$).

Table 4 shows firstly the estimated welfare statistics respectively based on the distributions of: y; x; x deflated with H; x deflated with $\bar{P}$; secondly, the simplified welfare statistics. The simplified statistics and the other estimated statistics are remarkably close. The differences are caused by non exact weak association of x and P. In quarter B, where the non rejection of independence is the strongest, the distance between true and simplified welfare statistics is very small. We have estimated bootstrap sampling confidence intervals. Because of the relatively small sample the confidence intervals (for the percentile bootstrap at 5% level) are much wider than the differences between true and simplified statistics. Since these significance results are uniforms and unambiguous, the estimated boundaries for these intervals are not shown in the table. Bootstrap

Table 4
Comparison of true, deflated and simplified statistics

		x/P	x/H	$x/\bar{P}$	x	Simplified statistics
σ	A	1701	1670	1646	1826	1711
	B	1620	1566	1547	1475	1605
	C	1968	1781	1750	1834	1836
	D	1511	1417	1406	1524	1437
CV	A	0.618	0.609	0.609	0.609	0.624
	B	0.599	0.581	0.581	0.581	0.595
	C	0.690	0.632	0.632	0.632	0.651
	D	0.654	0.617	0.617	0.617	0.626
Theil	A	0.166	0.158	0.158	0.158	0.165
	B	0.147	0.139	0.139	0.139	0.145
	C	0.192	0.168	0.168	0.168	0.177
	D	0.172	0.158	0.158	0.158	0.162
Mean	A	2750	2739	2701	2995	2739
	B	2702	2696	2663	2539	2696
	C	2850	2818	2770	2902	2818
	D	2310	2294	2275	2468	2294
$I_A^{0.5}$	A	0.0806	0.0765	0.0765	0.0765	0.0797
	B	0.0698	0.0657	0.0657	0.0657	0.0686
	C	0.0912	0.0809	0.0809	0.0809	0.0851
	D	0.0828	0.0769	0.0769	0.0769	0.0789
I_{GE}^2	A	0.190	0.185	0.185	0.185	0.195
	B	0.179	0.168	0.168	0.168	0.177
	C	0.237	0.198	0.198	0.198	0.212
	D	0.213	0.190	0.190	0.190	0.196

σ is the standard deviation. CV is the coefficient of variation. x is the nominal living standard. P is the Laspeyres price index. H is the harmonic mean of price indices. $\bar{P}$ is the arithmetic mean of price indices. The first (respectively second, respectively third, respectively fourth) column of numbers shows the value of the six calculated welfare statistics at every quarter (A, B, C, D), when using the distribution of x/P (respectively x/H, respectively $x/\bar{P}$, respectively x). The fifth column of numbers shows the value of the same welfare statistics when using the simplified formulae under weak association of x and P.

confidence intervals of the difference of the true and simplified statistics provide similar results.

The values obtained with the simplified formulae are almost always closer to the values based on y than that based on x. Moreover, deflating x with $\bar{P}$ and H and estimating the welfare statistics from these new distributions yield systematically worse results (or equivalent for the mean sometimes) than by using the simplified formulae. Besides, except in the case of the mean, using the distributions of x deflated by $\bar{P}$ or H does not even systematically deliver better results than those obtained by not deflating at all. Of course, for I_T, I_A, I_G and CV, which are scale-invariant, deflating by a constant price index has no effect and cannot improve the welfare estimate.

σ, CV, I_T, I_A, I_G are almost always slightly underestimated with the simplified formulae, but generally much less than with the statistics based on x only (or on x deflated by a constant price index). Although the correlation coefficients of $\ln x$ and $\ln P$ are not significant, they are not null (from -0.03 to -0.12, depending on the quarter) and affect the estimates. The fact that with these correlation levels the simplified formulae provide better results than the proposed alternatives, suggests that this may still occur with significant correlations.

6. Concluding remarks

We have investigated in this paper the spatial link of price indices and living standards. We first find that in rural Rwanda, the price index (P) is negatively related to the real living standard (y) at all quarters. This de facto spatial price discrimination against the poor is shown to be permanent and continuous across the whole living standard distribution. Various economic explanations are studied for this phenomenon, although none of them can be identified from the data. Further research is needed in this direction.

This result has several policy implications. First, direct analyses of living standard distributions should incorporate the association of y and P when dealing with policies changing the spatial distribution of prices. One expects that when the poor face relatively higher prices, a better efficacy of markets that would make spatial price indices less dispersed, would reduce poverty and inequality by removing the implicit price discrimination against the poor. There may then exist an opportunity for simultaneous enhancement of efficiency and equity similar to what is advocated by Bardhan (1996).

The second contribution of the paper is to show that the link of nominal living standards and Laspeyres (or Paasche) price indices is often likely to be attenuated and that when it is the case, simplified formulae of welfare indicators can be derived to permit the use of different information sources for prices and incomes. This also justifies the usual practice of separated study of the distributions of prices and incomes. The simplified formula for a given welfare indicator exhibits the few price statistics necessary for the correction of spatial price dispersion and that could be routinely be produced by statistical institutes. We show that these simplified formulae perform better in the case of rural Rwanda than the usual ad hoc price corrections. Note finally that even if the weak association may be rejected in other data sets, the correlation of the two variables may be small enough for the simplified formulae to be useful approximations.

Acknowledgements

I am grateful to the Ministry of Planning of Rwanda that provided me with the data, and in which I worked from 1984 to 1988 as a technical adviser from the French Cooperation and Development Ministry. I am also grateful to two referees and participants in the 2001 Meeting of the Economic Development Study Group in Nottingham for their comment. I am solely responsible for remaining errors.

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