COMPARING THE WEALTH OF NATIONS:
REFERENCE PRICES AND MULTILATERAL REAL INCOME INDEXES*

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ABSTRACT

This paper considers the problem of comparing real incomes across countries. The available methods are reviewed and their performance is compared using the raw data underlying the Penn World Table. The results throw light on the relative merits of different indexes and on the empirical importance of the "Gerschenkron Effect": the downward bias in a country’s measured real income when its own prices are used as weights. They also demonstrate the feasibility of using empirical demand parameters to estimate the GAIA ("Geary-Allen International Accounts") System.

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It is well known that international comparisons of real income are sensitive to the reference prices used. One hypothesis concerning the nature of the sensitivity involved is the so-called "Gerschenkron Effect": a country’s measured real income is higher the more the reference prices differ from its own prices. This effect is usually thought to reflect the bias arising from the lack of substitutability assumed in standard real income indexes. Less widely appreciated is the fact that even "true" indexes, which are based on utility theory and so allow for substitution, are in general sensitive to the reference price vector used. In this paper we set out to test the importance of these sensitivities, using the raw data which underlie the Penn World Table.

The other issue we address in the paper is the choice of index number for use in multilateral comparisons of real incomes between countries. Such comparisons have become increasingly common in economics in recent years, reflecting the greater availability of comparative data, of which the most extensive source is the Penn World Table which draws on the United Nations sponsored International Comparisons Project (ICP).1 Two principal methods are used in practice in such comparisons. The ICP and Penn World Table use a method originated by Geary (1958), which calculates world prices and exchange rates that corrected for deviations from purchasing power parity. However, it has been extensively criticised for its lack of theoretical foundations, notably by Samuelson and Swamy (1974) and Diewert (1981 and 1987). Diewert in particular has argued instead in favour of versions of the so-called EKS method, due to Elteto and Kojes (1964) and Szulc (1964), which is used

1 See Kravis (1984) and Summers and Heston (1991) for overviews of the ICP and the Penn World Table respectively.
by the OECD and by Eurostat.

Recent work by Neary (1996a, 1996b) argues that the claims made for the theoretical superiority of the EKS and related indexes do not hold up if tastes are not homothetic. Moreover, in that case, it is necessary to select from an infinity of "true" indexes to use as an ideal or benchmark. Neary proposes for such a benchmark the "GAIA" or "Geary-Allen International Accounts" System, which has the ease of interpretation of the Geary system but is consistent with utility theory in the context of multilateral comparisons. However, the ease of computation of such GAIA indexes and their relationship in practice with the Geary and EKS indexes remain open questions. These issues are addressed in the remainder of this paper. First, we present a brief overview of some relevant results from the theory of index numbers.\(^2\)

1. Index Number Theory

1.1 Fixed-Weight Bilateral Comparisons and the Gerschenkron Effect

Figures 1 to 3 illustrate the problems which arise in making bilateral comparisons of real incomes across countries. Figure 1 gives some hypothetical raw data: points J and K represent the quantities consumed while the slopes of the lines through those points represent the domestic relative prices of good \(1\) in each of two countries \(j\) and \(k\). It is convenient to choose one country as reference and, following ICP conventions, we select the high-income country \(k\). (This choice is arbitrary and does not affect the results in a substantive way.) This implies that we seek a measure of the real income of country \(j\) relative to that of country \(k\). It also implies that we measure expenditures in terms of country \(k\)'s currency, converting at current exchange rates. By this measuring rod, the two country's total expenditures in domestic prices equal OA and OK respectively.

The two countries' real incomes can now be compared at either set of domestic prices. Using country \(k\)'s prices leads to the Laspeyres measure, \(p_j^q/p_j^q\), given by OB/OK in Figure 2.\(^3\) By contrast, using country \(j\)'s prices leads to the Paasche measure, \(p_j^q/p_j^q\), given by OA/OK in Figure 3. The latter is clearly much lower, reflecting the "Gerschenkron Effect" due to Gerschenkron (1951). To quote Samuelson, "It is always better to ride the other man's horse"; more precisely, each country's relative real income is lower when the comparison is made at its own prices and higher when it is made at the other's. However, as we shall see, this outcome is not inevitable on theoretical grounds and so its empirical relevance is an open question which deserves investigation.

1.2 Reference Prices and True Bilateral Comparisons

The fact that fixed-weight indexes of real income are sensitive to the reference prices used suggests that we should consider a "true" or utility-based index instead. The most natural of these is the Allen (1949) quantity index, which equals the ratio of the expenditure functions of the two countries evaluated at a common reference price vector \(p'\):

\[
e(p', u) \quad (1)
\]

Since the expenditure function gives the minimum cost of attaining a given utility level facing given prices, this index allows for intercommodity substitution and so avoids the biases of fixed-weight indexes. However, it is not independent of reference prices, except in the special case of homothetic tastes. The expenditure function then takes the form \(e(p', u) = u.e(p')\), and so (1) reduces to the ratio of utilities, \(u/u'd\), which is independent of \(p'\). When

\(^2\) See Diewert (1981, 1987), Pollak (1971) and Neary (1996a) for more detailed accounts.

\(^3\) With \(m\) commodities, we let \(p^m\) and \(q^m\) denote the \(m\)-by-\(1\) vectors of prices and quantities respectively in country \(h\). A dot denotes a vector inner product.
tastes are not homothetic, the realistic case, the Allen index is sensitive to the choice of reference prices. Figures 4 and 5 illustrate the difference between what may be called "Laspeyres-Allen" and "Paasche-Allen" indexes, which use country \( k \) and country \( j \) prices as reference respectively.

While the Allen index does not (except when tastes are homothetic) avoid the dependence of our measure of real income on the reference prices used, we can use it to throw light on the Gerschenkron Effect. Since the expenditure function gives the minimum cost of attaining a given utility level facing particular prices, it follows that \( e(p^i, u^i) \leq p^i q^i \). Hence the Laspeyres-Allen index cannot exceed the corresponding fixed-weight Laspeyres index:

\[
\frac{e(p^i, u^i)}{e(p^k, u^k)} \leq \frac{p^i q^i}{p^k q^k}. \tag{2}
\]

By similar reasoning, the Paasche-Allen index cannot be less than the corresponding fixed-weight Paasche index:

\[
\frac{e(p^i, u^i)}{e(p^i, u^k)} \geq \frac{p^i q^i}{p^i q^k}. \tag{3}
\]

These two inequalities can be combined in the special case of homothetic tastes. Substituting for the special form of the expenditure function in this case, the left-hand sides of (2) and (3) are equal to one another and the Laspeyres index necessarily exceeds the Paasche index. (More precisely, the Laspeyres index cannot be less than the Paasche index and must strictly exceed it if any inter-commodity substitution occurs.) This gives a well-known result: the Gerschenkron Effect is a necessary consequence of utility maximisation by a single consumer whose tastes are homothetic.

Figure 6 illustrates the result. Points \( J \) and \( K \) represent countries \( j \) and \( k \) as before. Since tastes are homothetic, all indifference curves have the same slope along the ray \( O K \). In particular, the slope of the indifference curve through \( J \) at point \( K' \) is the same as the slope of the price line at \( K \). The Laspeyres-Allen index therefore equals \( O K'/O K \). It follows immediately that the fixed-weight Laspeyres index, which equals \( O B/O K \), must exceed the fixed-weight Paasche index, which equals \( O A/O K \).

However, a similar geometric approach may be used to show that, if tastes are not homothetic, the Gerschenkron Effect may not arise. In Figure 7 good 1 is income-elastic but nevertheless has a higher relative price in \( k \) than in \( j \). As a result, the Laspeyres index \( O B/O K \) is less than the Paasche index \( O A/O K \). The Gerschenkron Effect is violated, even though the data relate to a single utility-maximising individual. We may conclude from these examples that substitutability tends to encourage the Gerschenkron Effect and that non-homotheticity may, but need not, work against it. Of course, when we compare data for whole countries these results derived from individual behaviour need not apply.

1.3 Multilateral Comparisons

So far, we have considered only bilateral international comparisons. To compare the real incomes of a group of countries the issues already considered are still relevant and many others also arise. Rather than giving a full account, we summarise the three multilateral indexes whose empirical performance is considered in the remainder of the paper.

The Geary System: The method proposed by Geary (1958) postulates the existence of "world" prices \( \pi \) and "true" exchange rates \( \epsilon \). The true exchange rates are Laspeyres price

\[ \text{\textsuperscript{4}} \text{Nuxoll (1994) states that the Gerschenkron Effect is implied by the Weak Axiom of Revealed Preference when tastes are homothetic. For bilateral comparisons, the Weak Axiom is equivalent to utility maximisation.} \]
indexes, which compare the world prices with the prices of each country in turn:

\[ e_j = \frac{\sum_i \pi_i q_{ij}}{\sum_i \sum_j p_{ij} q_{ij}}, \quad j = 1, \ldots, m. \]  

(4)

As for the world prices themselves, they satisfy the property that total world spending on commodity ı is the same whether valued at its world price or at domestic prices converted at the true exchange rates:

\[ e_i = \frac{\sum_j e_j p_{ij} q_{ij}}{\sum_j q_{ij}}, \quad i = 1, \ldots, n. \]  

(5)

Solving for \( e_\) and \( \pi_\), it is then straightforward to calculate the income of each country at world prices:

\[ z_j^0 = e_j z_j = \sum_i \pi_i q_{ij}, \quad i = 1, \ldots, m. \]  

(6)

These real income measures in turn imply a set of indexes, \( Q_{jk}^0 = z_j^0 / z_k^0, \forall j, k. \) Thus the Geary method yields fixed-weight index numbers of real income which use as reference the prices \( \pi_\) of a hypothetical "world" country.

The EKS Index: The starting point for the EKS system is the Fisher Ideal index, which is the geometric mean of the base-weighted Laspeyres index and the current-weighted Paasche index:

\[ \ln Q_{jk}^0 = \frac{1}{2} \left[ \ln \frac{p^k q^j}{p^k q^k} + \ln \frac{p^j q^j}{p^j q^k} \right]. \]  

(7)

The Fisher Ideal index has many desirable properties but it is not suited to multilateral comparisons. The EKS index extends it to the multilateral context since it equals the geometric mean of the ratios of all \( m \) bilateral Fisher Ideal indexes, taking each of the \( m \) countries in turn as base:

\[ \ln Q_{jk}^{EKS} = \ln Q_{jk}^0 = \frac{1}{2} \sum_{i=1}^{m} \left\{ \ln Q_{ji}^0 - \ln Q_{ij}^0 \right\}. \]  

(8)

This index, unlike the Fisher index, yields a transitive ranking across countries and is not sensitive to the choice of base country. It also reduces to the Fisher index when \( m=2 \). Thus the EKS index is indeed an appropriate multilateral generalisation of the Fisher Ideal. Note that it is not possible to give a reference price interpretation to the EKS index.

The GAIA ("Geary-Allen International Accounts") System: This is an "ideal" counterpart to the Geary system, proposed by Neary (1996a and 1996b). Its starting point is to replace the fixed-weight Laspeyres formula in the Geary exchange rates (4) with their true equivalents, which may be called Geary-Konüs exchange rates:

\[ E_j = \frac{e(P, u^j)}{e(p^j, u^j)} = \frac{\sum_i \pi_i q_{ij}^*}{\sum_i p_{ij} q_{ij}}, \quad j = 1, \ldots, m. \]  

(9)

Here the \( q_{ij}^* \) denote the "virtual" or imputed quantities which country \( j \) would choose if it were faced with world prices \( P \):

\[ q_{ij}^* = e(P, u^j). \]  

(10)

As for the world prices \( P \), they must satisfy aggregation conditions of the Geary type. They cannot do so in terms of actual quantities consumed but they can in terms of imputed quantities. This leads to a set of Geary-Konüs world prices:
Finally, the implied Geary-Allen measures of income at world prices are defined as follows:

\[ z_j^* = \frac{E_j \Pi_j}{\sum_i \Pi_i}, \quad j = 1, \ldots, m. \]  

(12)

As noted in Neary (1996b), this system combines the theoretical consistency of Allen indexes with the ease of interpretation of the Geary system. It also allows an interpretation of the world prices: if preferences are characterised by "generalised linearity" as proposed by Muellbauer (1975), they correspond to the prices of a hypothetical country whose income is an appropriate average of world incomes. However, notwithstanding these theoretical advantages, the GAIA system is unobservable and so its empirical relevance depends on how easily and plausibly it can be estimated using available information on demand parameters.

We investigate this issue in Section 4 below. First we consider the sensitivity of both fixed-weight and true indexes to the choice of reference prices.

2. Laspeyres Indexes and the Gerschenkron Effect

In this section and the next we examine empirically the significance of the issues raised in the theoretical discussion above. The data we use are taken from the International Comparison Project (ICP) which underlies the Penn World Table and are described in detail in the Appendix. In this pilot study we use data on 16 countries in 1970, giving the prices and quantities consumed of 11 categories of personal consumption. All prices are measured in current U.S. dollars, converted at market exchange rates.

Table A1 in the Appendix gives the raw data. The first step is to calculate the income of each country relative to the U.S., using the prices of each country in turn as reference, and the results of this are given in Table 1. The countries are ranked by \( z_j \), their per capita expenditure in dollars, converted at current exchange rates. For reference, this is given in levels and relative to the U.S. in the first two rows of the table. The next sixteen rows, constituting the bulk of the table, give the relative Laspeyres indexes of real income, using the prices of each country in turn as reference. Reading down each of the country columns gives alternative estimates of the real income of each country relative to the U.S. Reading across each row gives the indexes corresponding to a particular country's prices; i.e., the incomes of each country relative to the U.S. using a particular country's prices as weights. Thus the entry in row \( k \) and column \( j \) equals \( p^k q^k / p^j q^j \) over \( p^k q^k / p^j q^j \); i.e., \( p^k q^k / p^j q^j \), the real income of country \( j \) relative to country \( l \) (the U.S.) using the prices of country \( k \) as reference. This matrix may be called the Laspeyres "star" matrix, since each row gives a star index where the reference country is the centre and the other countries are the points of a star. The next three rows and the final three columns of the table give the average, standard deviation and coefficient of variation of the different Laspeyres indexes in each row and column. The final two rows of the table give the values of the EKS and Geary indexes of real expenditure.

The first major feature of the Laspeyres star matrix is that all the different indexes show less dispersion in real incomes than the data on nominal expenditure. Thus the coefficient of variation of nominal expenditure across the sixteen countries is 98.2%, whereas the coefficient of variation of the Laspeyres indexes ranges from 81.3% to 85.8%, with an average across countries of 83.5%.\(^{5}\) The different indexes also give broadly similar

\(^{5}\) These coefficients of variation are independent of the choice of base country. With the U.S. as base, the reduction in dispersion shows up as a higher average real income with a near-identical standard deviation.
rankings across countries which may differ from the rankings based on nominal expenditure. For example, almost all the different indexes of real expenditure rank Germany below each of Belgium, France and the Netherlands, and rank Korea below Iran, in all cases reversing the corresponding rankings of nominal expenditure.

While there are major similarities between the indexes there are also important differences. These are highlighted by the summary statistics below the star matrix, in particular by the row giving the coefficient of variation of the indexes for each country. The average of these for the fifteen countries excluding the U.S. is 5.4%, indicating a significant variation in the estimate of real income depending on the choice of weights. This effect is considerably greater for poorer countries: the correlation coefficient (over fifteen countries) between the coefficient of variation and the level of nominal expenditure is −0.824.

We next wish to examine the magnitude of the Gerschenkron Effect in the sample. This means considering all possible bilateral comparisons. With m countries there are m(m−1)/2 such comparisons; 120 in our 1970 sample of 16 countries. For each of these comparisons we treat the ratio of the Laspeyres to the Paasche index, $p'_i q_i/p'_i q'_i$, divided by $p'_i q'_i/p'_i q'_i$, as an independent observation and we examine the sample distribution of these observations. The results are given in the first row of Table 2.

Considering first the properties of the distribution, the table shows that the average of the 120 ratios was 1.114, with a standard deviation of 0.118. Although this is not significantly greater than unity, the distribution is highly skewed, with 112, or 93.3%, of the ratios exceeding unity. This suggests that the Laspeyres index is indeed systematically greater than the corresponding Paasche index. A more formal test of the Gerschenkron Effect requires that we investigate how their ratio varies with the difference in prices between the two countries. To do this we estimate a simple regression equation which regresses the Laspeyres to Paasche ratio on the sum of squared deviations between prices of individual goods in the two countries; i.e., a measure of the extent of price dispersion between the two countries. As the table shows, the Gerschenkron Effect is overwhelmingly confirmed. A unit increase in this measure of price dispersion is significantly associated with a rise in the Laspeyres to Paasche ratio of 0.04704.

3. Allen Indexes Based on Estimated Stone-Geary Preferences

The previous section has confirmed that the choice of reference prices makes a significant difference to the estimated levels of real expenditure and that the differences involved reflect the Gerschenkron Effect. This suggests the desirability of estimating alternative indexes which allow for substitutability between commodities in response to price differences. Of course, there are many ways in which such substitutability might be parameterised. We have chosen to do so in a simple fashion, using the linear expenditure system, which corresponds to the Stone-Geary utility function. The demand function for commodity group $i$ in this system (suppressing country subscripts for simplicity) is:

$$p_i q_i = p_i y_i + \beta_i (z - \sum_k p_k y_k)$$

(13)

The $\gamma_i$ coefficients are usually interpreted as subsistence parameters, while the marginal budget shares $\beta_i$ (which must sum to unity) determine the allocation of "supernumerary" income $z - \sum_k p_k y_k$ among commodity groups.

Table 3 gives the parameters of the linear expenditure system estimated from the 1970 data for 16 countries. The statistical performance of the equations is satisfactory and the

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4 In this pilot study we did not experiment with alternative specifications. The estimation method is Seemingly Unrelated Regressions, with no correction for heteroscedasticity.
see that food, beverages and tobacco decline in importance as total expenditure rises, whereas
gross rents, fuel and power, medical care, transport and communications, recreation and
education and miscellaneous items are luxuries.

Using these results to calculate Allen indexes is straightforward. The expenditure
function corresponding to the Stone-Geary utility function, evaluated at the prices of country
k and the utility level of country j is:

\[ e(p^k, u^j) = \gamma^k - \beta^k u^j, \quad \text{where: } \gamma^k = \sum_i u_i p_i^k \quad \text{and} \quad \beta^k = \prod_i \left( \frac{p_i^k}{\bar{p}_i} \right)^{\frac{1}{k}}. \] (14)

The \( \gamma^k \) and \( \beta^k \) terms can be calculated using the parameter estimates from Table 3, while the
level of utility in each country is found by solving (14) for \( j = k \) (when the left-hand side
equals \( z' \) and so all the terms in the equation except \( u \) are observable). All 16-by-16 terms
implied by (14) can then be calculated and the results are shown in the central matrix in
Table 4. The entry in row \( k \) and column \( j \) of the matrix gives the real income of country \( j 
relative to the reference country (the U.S.), evaluated at the prices of country \( k \); that is,
\[ e(p^k, u^j)/e(p^j, u^j) \text{ or } e(p^k, u^j)/e(p^j, u^j). \] The special demand system used
imposes considerable structure on the entries in the matrix. From (14), the difference
between any two entries \( j \) and \( h \) (not necessarily adjacent) in row \( k \) is:

\[ \Delta^h = \frac{u_j - u_h}{u_k^j} H_k, \quad \text{where: } H_k = \frac{u_k^j \beta^k}{\gamma^j + u_k^j \beta^k}. \] (15)

In words, this is the utility difference between countries \( j \) and \( h \) (relative to the U.S.) times
\( H_k \), which measures the ratio of supernumerary to total income for the U.S. at country \( k 
prices. Thus, any entry in the table can be calculated from the values of \( u \) and \( H \) only (given

prices. Thus, any entry in the table can be calculated from the values of \( u \) and \( H \) only (given
in the first row and last column respectively).\(^8\)

Since all the indexes are linear in utility, they are perfectly correlated across columns.
Nevertheless, reading down columns there is considerable variation between the different
indexes for each country. This is less true of the high-income countries in the sample, for
which the results are relatively insensitive to the reference prices used. However, for the
low-income countries the opposite is true. In fact, the average coefficient of variation (over
fifteen countries) for these indexes is 10.4%, almost twice as much as for the Laspeyres
indexes in Table 1.\(^9\) Of course, some of this variation may reflect the failure of the linear
expenditure system to capture the variations in expenditure patterns between the countries in
the sample.\(^10\) Nevertheless it suggests that the choice of reference prices makes at least as
much difference to the true indexes as it does to the fixed-weight Laspeyres indexes.

Finally, applying to the Allen indexes the same tests for the Gerschenkron Effect as we
applied to the fixed-weight indexes in Section 2, we find that the presence of the effect is
rejected. (See the second row of Table 2.) This shows that the sensitivity of the Allen

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\(^8\) Equation (14) also applies to any member of the Gorman polar form family of demand
systems, which generalise the linear expenditure system to allow for arbitrary \( \gamma^k \) and \( \beta^k 
functions which are linearly homogeneous in prices. Hence the comments in this paragraph
apply to any member of this family.

\(^9\) As with the Laspeyres indexes, this effect is greater for poorer countries though rather less
so than in Table 1: the correlation coefficient (over fifteen countries) between the coefficient
of variation and the level of nominal expenditure is \(-0.532\).

\(^10\) The two negative utility levels for India and Kenya pose special difficulties of
interpretation. Since utility is ordinal rather than cardinal, negative values do not matter in
principle, but with Stone-Geary preferences they imply that the country in question has total
expenditure less than the *subsistence* level (i.e., \( z' < \sum p_i \gamma_i \)). This in turn implies that
the substitution matrix is not negative semi-definite when evaluated at those countries'
consumption bundles. This finding is common with the linear expenditure system. Brown
and Deaton in their 1972 survey (Section IV.3) note that even in time-series studies
supernumerary income is often negative for the first few observations.

\(^7\) Our discussion in the remainder of this section has benefitted from a stimulating
correspondence with Patrick Honohan.
between the countries compared.

4. Groping Towards GAIA

The last set of indexes we calculate are those corresponding to the GAIA system. To do this we implement the iteration procedure proposed in Neary (1996b). This is a tâtonnement-type algorithm, which takes as starting point the standard Geary world prices and true exchange rates and uses them to calculate initial estimates of the virtual quantities defined in (10). The Geary method is then applied to the estimated virtual quantities and the process is continued until it converges.

In practice, it proved possible to implement this procedure fairly straightforwardly. The convergence criterion selected was that the sum of squared deviations between the estimated \( \bar{q} \) vectors in successive iterations be less than \( 10^{-14} \). The results are given in Tables 5 and 6. The convergence criterion was satisfied after 316 iterations, though to the degree of precision given in the tables there was no difference between iterations 251 and 316. The tables show that the procedure converges reasonably rapidly: very much so for real incomes, rather less so for exchange rates and slowest of all for world prices.

Table 7 summarises the relationships between the principal indexes which we have calculated as well as some of the statistical properties of each. As can be seen the correlations between all the indexes are extremely high. (The average of the Allen indexes is perfectly correlated with the GAIA index, reflecting the fact that, with Stone-Geary preferences, both these indexes are linear functions of the data.) The expenditure index is the least correlated with the other indexes while the GAIA index is slightly more correlated with the EKS index than with the Geary index. Of course, the high correlations between the different indexes does not mean that they have identical implications for the levels of real consumption expenditure in different countries. Note finally that the GAIA index is the most dispersed of all the indexes, except for that based on nominal expenditure.

5. Conclusions

This paper has used a specially constructed data set drawn from the International Comparisons Project which underlie the Penn World Table to investigate a number of related issues which arise in making comparisons of real incomes and expenditures between countries.

The first issue we addressed was the empirical importance of the Gerschenkron Effect, which postulates that a country's relative real income is higher the more the reference prices used to make the comparison deviate from its own prices. We noted that, on theoretical grounds, this phenomenon may be expected to occur to the extent that utility-maximising consumers substitute away from commodities which are relatively more expensive. However, when tastes are not homothetic and income levels are sufficiently different, it may not occur even if we had observations on the behaviour of individual utility-maximisers.

Given these theoretical ambiguities concerning the likelihood of the Gerschenkron Effect, we proposed and implemented two tests of its severity, based on the distribution of the ratios of the Laspeyres to Paasche indexes for all possible bilateral comparisons in the sample. Our results confirmed that the Gerschenkron Effect was present in the sample. 93.3\% of the ratios exceeded unity and they were significantly and positively related to the degree of price dispersion between the two countries compared. The same tests were applied to the Allen indexes, calculated using the estimated parameters of the Stone-Geary utility function. By

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11 Copies of the GAUSS program which calculates the GAIA system are available on request from the authors. Implementation on a Pentium 90 notebook with 16MB RAM running under Windows 95 took 2.69 seconds of calculation time.
contrast with the results for the Laspeyres indexes, there was no significant relationship
between a country’s Allen real income index and the deviation between its own prices and
the reference prices used. This suggests that taking account of intercommodity substitution
substantially eliminates the Gerschenkron Effect.

While the Allen indexes eliminate the Gerschenkron Effect, they do not reduce the
dependence of real income measures on reference prices. On the contrary, a comparison of
Tables 1 and 4 shows that the sensitivity of real incomes to reference prices is considerably
greater for the Allen than for the Laspeyres indexes, with the increased sensitivity especially
pronounced for low-income countries. This strengthens the case for a system such as the
GAIA which provides an explicit rationale for the choice of reference prices.

Finally, the actual estimation of the GAIA system was shown to be feasible, with
convergence occurring rapidly. The resulting indexes are more dispersed than any of the
other real income indexes, though they are slightly more strongly correlated with the EKS
than with the Geary indexes.

Naturally, this paper has raised more questions than it has answered. It would be
especially desirable to extend the methods of calculating real expenditures to other categories
of aggregate demand, with a view to calculating true indexes of national output. As for the
tue indexes which we have estimated, they are clearly conditional on the hypothesis that
demands in all the countries of the sample can be represented by the linear expenditure
system. Since the Stone-Geary utility function which underlies this exhibits additive
separability, it is open to the criticism of Deaton (1974) that it imposes on the results an
inverse relationship between income and price elasticities. (Deaton calls this “Pigou’s Law”.)
This may explain why the EKS system does not fare badly relative to the GAIA system;
deviations from homotheticity work against the EKS system but, because of Pigou’s Law,
they also tend to impose greater price substitutability, which also works against the Geary
system. These considerations suggest that it would be very desirable to repeat the
calculations of this paper using demand systems which do not impose additive separability.

Appendix: The Data

The results in this paper use data for 1970 only. However, it is convenient to describe
the full data set which has been assembled. The raw data are taken from Phase II (1970),
Phase III (1975), Phase IV (1980) and Phase V (1985) of the United Nations International
Comparisons Project (ICP). Phase IV is available in hard copy as United Nations (1986)
while the other phases were retrieved, via the Internet, from files at the University of
Pennsylvania. These sources give data on 16 countries for 1970, 34 countries for 1975, 60
countries for 1980 and 64 countries for 1985.

The present project uses 11 categories of personal consumption expenditure: food;
beverages; tobacco; clothing and footwear; gross rents; fuel and power; house furnishings,
appliances and operations; medical care; transport and communication; recreation and
education; and miscellaneous goods and services. The data provided for 1980 were already
aggregated to this level whereas data for the other three years were not.

To calculate the data for 1980, Tables 6 and 8 of United Nations (1986) are used. Table
6 provides data on per capita expenditure in national currencies, \( z_i = p_i q_i \) where \( p_i \) is the
price of good \( i \) in country \( j \) and \( q_i \) is the quantity of good \( i \) in country \( j \). Table 8 provides
the purchasing power parities which are the national currency expenditures from Table 6
divided by expenditure in international prices. These international prices are produced by
the Geary method of aggregation used in the ICP. Therefore the entries in Table 8 are:
\[
\frac{p_i q_i}{\pi_i q_i} = \frac{p_i}{\pi}, \quad \text{where} \quad \pi_i \text{ is the international price of good } i.
\]
by the corresponding entry for the United States, \( p_j / \pi_j \), gives prices in country \( j \) relative to prices in the United States: \( p_j / p_d \). Dividing each entry in Table 6, \( p_j / p_d \), by the corresponding relative price, \( p_j / p_d \), gives quantities in country \( j \) measured in U.S. prices, \( p_j / q_j \). This gives \( p / p_d \) and \( p_d / q_j \), which are the price and quantity data required to calculate the various real income indexes.

For 1970, 1975 and 1985, a different procedure was followed. Expenditure data for these years were given as percentages of GDP while the price data were already given in the desired form relative to the US, \( p_j / p_d \). The data were broken down into 150 categories on average, of which 110 relate to consumption, so these had to be aggregated to obtain our 11 categories. To get expenditure on food in country \( j \), for example, the expenditure shares of all items in the food category were summed and then grossed up by GDP, taken from the IMF statistical yearbooks. The aggregate price level for food was calculated as the weighted sum of all food item prices where the weight was the share of each food item in total food expenditure. The same procedure was then followed to derive price and quantity data from the price and expenditure data for each of these years as was used to derive the 1980 data. By way of illustration, Table A1 gives the calculated raw data for 1970.

References


comparisons of purchasing power and real incomes,* Economic and Social Review, 27:2 (January), 161-179.


Szulc, B.J. (1964): "Index numbers for multilateral regional comparisons" [in Polish], Przegląd Statystyczny, 3, 239-254.


Table 1: Laspeyres, EKS and Geary Indexes of Real Consumption Expenditure per capita, 1970

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S.D.: Standard Deviation; C.V.: Coefficient of Variation (%)
### Table 2: Tests of the Gerschenkron Effect

<table>
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<tr>
<th>Index</th>
<th>Distribution of Laspeyres to Paasche Ratios</th>
<th>Regression of LP Ratio on PD</th>
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<td>1.528</td>
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All comparisons are based on the 120 ratios of bilateral Laspeyres to Paasche indexes for the 16 countries in 1970

PD: Measure of price dispersion, equal to the sum of squared deviations of prices between the two countries

Regression coefficient estimates have standard estimates in parentheses

### Table 3: Estimated Parameters for the Stone-Geary Utility Function, 1970

<table>
<thead>
<tr>
<th>Commodity Group</th>
<th>Average Budget Shares</th>
<th>β's: Marginal Budget Shares</th>
<th>γ's: Subsistence Parameters</th>
<th>R squared</th>
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<td>t-value</td>
<td>Coeff.</td>
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<td>Miscellaneous</td>
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* β for the Miscellaneous group was estimated residually, so these entries are not relevant
Table 4: Allen Indexes of Real Consumption Expenditure per capita, 1970

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<th>Country</th>
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<th>France</th>
<th>Netherlands</th>
<th>UK</th>
<th>Italy</th>
<th>Japan</th>
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<th>Malaysia</th>
<th>Philippines</th>
<th>India</th>
<th>Kenya</th>
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<td>C.V. (%)</td>
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<td>H: Ratio of Supernumerary to Total Income for U.S. at Country k Prices</td>
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<td>7.5</td>
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</table>
| S.D.: Standard Deviation; C.V.: Coefficient of Variation (%)

Table 5: Convergence Towards G A L A - Real Incomes and True Exchange Rates

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<th>UK</th>
<th>Italy</th>
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</tr>
</tbody>
</table>
| S.D.: Standard Deviation; C.V.: Coefficient of Variation (%)

* S.D.: Standard Deviation; C.V.: Coefficient of Variation (%)

** Exceptionally, these entries are based on fifteen countries only (excluding the U.S.)
### Table 6: Convergence Towards GAIA - World Prices

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### Table 7: Correlations between Different Indexes

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<th>Correlation Coefficients</th>
<th>Expenditure</th>
<th>Av. Laspeyres</th>
<th>Av. Allen</th>
<th>EKS</th>
<th>Geary</th>
<th>GAIA</th>
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<td>0.9966830510</td>
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Mean: 287.3 337.2 334.5 332.6 337.2 331.5
Standard Deviation: 282.2 281.5 281.6 281.4 280.9 282.9
Coefficient of Var (%): 98.2 83.5 84.2 84.6 83.3 85.3
### Table A1: Raw Data 1970

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| **Quantities** |      |       |      |        |       |     |      |       |       |        |       |       |        |       |       |       |
| Food           | 470.94 | 268.63 | 385.76 | 396.59 | 357.15 | 365.23 | 397.43 | 287.36 | 87.73 | 144.50 | 149.43 | 129.25 | 105.71 | 123.61 | 79.32 | 3.55 |
| Beverages      | 68.02  | 65.01  | 56.70  | 101.20 | 74.34  | 82.27  | 61.28  | 70.93  | 18.17 | 15.58  | 6.71   | 1.37   | 6.02   | 13.23  | 0.74  | 0.10 |
| Tobacco        | 55.50  | 22.82  | 53.25  | 33.42  | 54.54  | 47.09  | 32.96  | 24.59  | 21.30 | 9.74   | 7.95   | 10.84  | 8.51   | 4.01   | 0.09  |
| Clothing       | 251.06 | 159.23 | 134.69 | 120.62 | 174.45 | 142.46 | 108.73 | 140.65 | 41.60 | 41.09  | 29.01  | 31.19  | 22.24  | 16.66  | 7.54  | 0.27 |
| Gross rents    | 485.26 | 255.62 | 286.44 | 239.49 | 120.37 | 220.70 | 145.76 | 27.03  | 29.06 | 7.45   | 119.00 | 50.92  | 53.02  | 13.05  | 0.49  |
| Fuel & power   | 97.16  | 35.82  | 36.74  | 28.83  | 60.03  | 38.42  | 22.71  | 15.23  | 4.64  | 5.91   | 1.97   | 5.85   | 2.46   | 1.89   | 0.61  | 0.01 |
| House          | 262.03 | 210.87 | 236.33 | 139.77 | 206.18 | 123.50 | 74.67  | 109.65 | 38.85 | 25.39  | 21.02  | 21.57  | 15.97  | 8.08   | 0.55  |
| Medical care   | 207.40 | 190.60 | 90.54  | 209.94 | 167.36 | 80.35  | 156.81 | 218.00 | 65.45 | 25.82  | 16.47  | 21.11  | 17.50  | 2.63   | 4.13  | 0.48 |
| Transport      | 467.61 | 164.03 | 137.77 | 137.89 | 126.76 | 172.72 | 105.87 | 58.22  | 18.94 | 43.49  | 29.92  | 11.86  | 31.33  | 5.75   | 9.11  | 0.34 |
| Recreation     | 470.90 | 222.05 | 237.92 | 190.94 | 217.46 | 260.03 | 167.99 | 141.85 | 67.06 | 35.23  | 35.42  | 20.98  | 49.33  | 28.26  | 13.55 | 0.83 |
| Misc.          | 426.23 | 207.12 | 379.78 | 238.39 | 261.85 | 171.35 | 98.08  | 300.33 | 85.78 | 70.67  | 31.90  | 77.50  | 23.50  | 34.59  | 8.37  | 0.66 |
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