# Substitution Biases in Price Indexes during Transition.

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## Abstract

The rapidly changing environment of the transition may create special problems for calculation of index numbers that require a fixed basket of goods and retail outlets. Using referent-level data we find that fixed-weight Laspeyres index on average overstated cost of living increases by approximately 5 per cent a year when compared with a superlative index in the Czech Republic. This difference is smaller than might be expected given the large changes in relative prices that occurred during transition and suggests that consumer substitution impacts may have been largely offset by other factors, especially rising prices combined with increased consumption of some goods as artificial shortages under communism were removed. Indeed, in the period of greatest supply response to price liberalization, the Laspeyres index appears to understate increases in the cost of living.

JEL:    C82, E31, P24

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## I.	Introduction and Motivation

The issue of CPI measurement is particularly important for transition economies. Post-communist countries have experienced both much greater average inflation levels and larger shifts in consumption patterns than stable economies such as the U.S. Moreover, substantial variation exists across the region with respect to the frequency of revision in the basket used to measure prices changes. Moving from a planned, shortage economy to a market economy has involved massive changes in the availability and relative prices of various products. Thus, in the transition environment, not only will classic problems of CPI measurement be exacerbated, but other problems not commonly seen in more established economies may appear, in particular due to the relaxation of shortage constraints at a time of changing relative prices. As Triplett (2001, page 316) observed, even in relatively stable economies: "Politicians and policy makers are often subsequently surprised when price effects actually influence consumers' behaviour in the ways not considered when policies were adopted." In the rapidly changing transition environment it might be expected that behavioral responses would pose even greater challenges.

Despite these difficulties, transition economies require accurate measurement of price changes. All countries in the region must restrain government revenues and tax rates in order to promote economic growth, despite political pressures for large social programs indexed to inflation. If inflation is biased upward, therefore, apparent income declines during the transition are overstated, which could have clear domestic political consequences, including lower public support for necessary reforms and possibly increased political instability. In addition,

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understating of real incomes will cause accession countries to appear poorer than they in reality are, thereby indicating a higher need for transfers from current EU members and increasing difficulty of accession negotiations.

In recent years several studies have investigated the accuracy of Consumer Price Indices (CPI) as true measures of inflation. Almost all U.S. studies have concluded that the reported CPI overstates inflation.[1] Similar results exist for the U.K. (Oulton, 1996), Sweden (Dahlen, 1994), Australia (Woolford, 1994), and Germany (Hoffmann, 1999), among the other countries. One key reason that CPI measures are generally assumed to overstate the economically meaningful concept of a cost of living index is the ability of consumers to substitute away from products whose relative price has increased. Several studies (for example, Manser and McDonald, 1988 and Aizcorbe and Jackman, 1993) have shown that a conventional Laspeyres index increases about 0.2 percentage points a year faster in the U.S. than superlative indices that are better theoretical measures of changes in costs of living. It is unclear how this difference is affected by inflation rate or time since the last revision of the market basket (Moulton, 1996).

Substitution effects should be especially important in transition economies which have seen massive changes in relative prices. Koen and de Masi (1997) report year-to-year price correlations across various items in the range of 0.7 to 0.9 (and occasionally as low as 0.4) in post-communist economies. This contrasts with year-to-year correlations across items of 0.99 in the United States consumer price basket. The existence of such massive changes in relative

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[1] Gordon (1995), Boskin et. al. (1995), Darby (1995) and Diewert (1995) estimated a CPI bias above one percentage point per year, while Shapiro and Wilcox (1996), Jorgenson (1995), and Griliches (1995) concluded that the bias is about one percentage point per year. On the other hand, Wynne and Sigalla (1994), Lebow, Roberts and Stockton (1994), Pakes (1995) and Congressional Budget Office(1995) presented estimates of a lower CPI bias.

prices has led many observers to surmise that Laspeyres price indices will overstate cost of living changes by even more in the transition context than they do in more stable economies.[2]

There may, however, be offsetting tendencies. The planned economies were shortage economies where consumer demand often went unsatisfied (Kornai, 1992). Unsatisfied demand is symptomatic of prices set below market clearing levels. In such cases, the economic transition may result in increased purchases of goods whose prices are rising as liberalization results in establishing market-clearing prices that call forth additional supply to eliminate shortages. In these circumstances, price indices that use base-period weights (Laspeyres indices) may understate true cost of living increases. Thus, the extent, and even the direction, of the discrepancy between price and cost of living indices is an empirical question in the transition environment.

We use data from the Czech Republic to analyse the relationship between inflation measures using base-period and current-period weights during the transition. After a brief discussion of theoretical concepts we present estimates of this difference that is of the same relative order of magnitude as that found in the United States. While it is possible that this pattern results from consumers not responding to the massive shifts in relative prices that occurred during the transition, we find such an explanation inherently unlikely. Thus, it appears probable that other factors have largely offset substitution effects as markets reach equilibrium during transition. Finally, we analyse the sensitivity of our results to the level of aggregation at which the analysis is performed.

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[2] For discussions of possible biases in inflation during transition, see Brada, King and Kutan (2000), Filer and Hanousek (2000), Koen and De Masi (1997), and Duchene and Gros (1994).

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## II.    Differences Between Consumer Price and Cost of Living Indices:

Theoretically, two different conceptual frameworks can be used to measure changes in prices.  The first assumes a fixed level of consumer utility and produces a *cost of living index* (CLI), defined as the ratio of minimal costs needed in the two periods to achieve a given standard of living.  The second assumes a fixed basket of goods and results in a *consumer price index* (CPI).  Although it would appear that cost of living is a more relevant concept for public policy purposes, in practice, statistical offices generally opt  to compute a price index, usually a modified Laspeyres index where items in the market basket are weighted according to consumption patterns in an initial base period.[3]  There are several reasons  for this choice.  A Laspeyres index has low data requirements, enabling timely and routine calculation.  It is also consistent at various levels of aggregation.[4]

Given a sample of *n* prices, the relative price change (Laspeyres index) between period *T* and *T-1* is given by:

$$L_{T,T-1} = \frac{\sum_{j=1}^{n} Q_{j,B} P_{j,T}}{\sum_{j=1}^{n} Q_{j,B} P_{j,T-1}} \qquad (1)$$

where:

$$Q_{j,B} = \frac{E_{j,B}}{P_{j,B}}. \qquad (2)$$

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[3]Triplett (2001) provides a good discussion of these two alternatives.

[4] In other words, results of the Laspeyres formula are not affected by the level of aggregation. Computations made at different levels and aggregated will yield the same result as if made at a higher level directly (see, for example, Eichhorn and Voeller, 1976).

As is standard, $P$ stands for prices, $Q$ for quantity, and $E$ for expenditures. Subscript $j$ indicates the good, while $T$, $T$-$1$ and $B$, denote the current, immediately past, and base period, respectively. Substituting equation (2) into equation (1) yields a modified[5] Laspeyres index:

$$L_{T,T-1} = \frac{\sum_{j=1}^{n} [E_{j,B}/P_{j,B}]P_{j,T}}{\sum_{j=1}^{n} [E_{j,B}/P_{j,B}]P_{j,T-1}} = \sum_{j=1}^{n} w_{j,B}\left(P_{j,T}/P_{j,T-1}\right) \tag{3}$$

where $w_{j,B}$ is a weight based on consumer expenditures during the base period.

Alternatively, if we use period $T$ weights, the resulting price index is a Paasche index:

$$P_{T,T-1} = \frac{\sum_{j=1}^{n} [E_{j,T}/P_{j,T}]P_{j,T}}{\sum_{j=1}^{n} [E_{j,T}/P_{j,T}]P_{j,T-1}} = \sum_{j=1}^{n} w_{j,C}\left(P_{j,T}/P_{j,T-1}\right), \tag{4}$$

where $w_{j,C}$ is a weight based on current period ($T$) consumer expenditures.

It is obvious from the definition of the CPI that, when consumer behaviour remains unchanged across time periods, (i.e., the elasticity of substitution between products and outlets is equal to zero), a price index is also a true cost-of-living index. Otherwise indices based the concepts of fixed consumer basket versus fixed consumer utility may differ substantially. This possible divergence has led to the use of *superlative* price indices that are based on quantities (or expenditures) for both the initial and current period (Diewert, 1976). Perhaps the most common of the superlative indices is the Fisher exact index, defined as the geometric mean of the Laspeyres and Paasche indices, i.e.,

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[5] The term "modified" is used to indicate that the base period used to determine weights is earlier than the starting period for measuring price changes.

$$F = F_{T,T-1} = \sqrt{L_{T,T-1}P_{T,T-1}}. \tag{5}$$

Diewert (1976) has shown that if the underlying utility function is homothetic, then superlative indices give a close approximation of the true CLI..[6]

It is commonly assumed that Laspeyres indices exceed the corresponding Paasche indices and that, therefore, Laspeyres price indices overstate increases in the cost of living, while Paasche indices underestimate increases in living costs. This widely spread assumption comes from the well-known fact that when prices are market clearing and the utility function is homothetic, then the true CLI lies between an upper bound determined by the Laspeyres index and a lower bound given by the Paasche index.[7] In other words, this intuition is based on an assumption that changes in consumer purchasing pattern are demand-side driven and occur because of responses to changes in relative prices. In this model, consumers respond to exogenous shifts in prices by substituting relatively less expensive alternatives for items whose price has increased. Thus, weighting the CPI basket according to initial period weights will overweight the products that exhibit increases in their relative prices, causing a Laspeyres Index to exceed a comparable Paasche Index. This conventional situation is illustrated in Example 1.

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[6] It is unlikely that this assumption is strictly satisfied in reality but such superlative indices are widely used. Their validity also relies on an assumption that consumer tastes remain unchanged during the periods studied (see Diewert, 1976).

[7] For a general discussion of various index numbers, see Fisher (1927), Frisch (1936), Adelman (1958), Eichorn and Voeller (1976), and Forsyth and Fowler (1981).

**Example 1**.    **Substitution in Response to Exogenous Supply Shift**

|  | Price in Period 1 | Quantity in Period 1 | Price in Period 2 | Quantity in Period 2 |
|---|---|---|---|---|
| Beef | 1 | 1 | 1.5 | 0.4 |
| Chicken | 1 | 1 | 1 | 1.4 |

**Laspeyres index     1.25           Paasche index          1.11**

There are, however, scenarios other than that of substitution caused by exogenous relative price shifts.  Suppose, for example, that the relative price changes are endogenous because a taste shift has caused consumer interest in a particular product to change, thereby causing demand and price for that product to shift in the same direction.  Examples might include a scare related to BSE causing demand for beef to plummet, or research linking cholesterol to heart disease reducing demand for eggs despite a decline in their relative price (see Brown and Schrader, 1990).  Alternatively, and especially relevant to the transition case, price controls may result in prices below equilibrium levels while simultaneously depressing supply (see Osband, 1992, who discusses biases under conditions of shortage in the starting period).  In such a case, removal of controls might result in the outcome shown in Example 2.

**Example 2.    Demand Shift Due to Taste Changes or Removal of Price Controls**

|  | Price in Period 1 | Quantity in Period 1 | Price in Period 2 | Quantity in Period 2 |
|---|---|---|---|---|
| Beef | 1 | 1 | 0.5 | 0.4 |
| Chicken | 1 | 1 | 1.2 | 1.5 |

**Laspeyres index     0.85           Paasche index          1.05**

Here, the Paasche index exceeds the Laspeyres index. Thus, the direction of the bias created by using base-period weights to calculate consumer price indices in the transition environment is an empirical question. The existence of large shifts in relative prices may create greater than normal substitution incentives, leading to a large overstatement of the true cost of living increase. On the other hand, removal of disequilibria may cause Laspeyres indices to actually understate true inflation.

### III.     Comparison of Consumer Price and Cost of Living Indices for the Czech Republic

In this section we compute Fisher exact indices for various years for the Czech Republic and compare them to the official modified Laspeyres indices used to measure consumer price inflation for the same years. We focus on the Czech Republic for several reasons. First of all, revision of the consumer basket occurred only every five years (a frequency similar to the U.S. but substantially less often than in other transition countries such as Hungary and Bulgaria). Secondly, the methodology used is similar to that of the Bureau of Labor Statistics (BLS). Finally, the Czech Statistical Office (CSO) provides a very detailed description of its methods, including documentation of the price and expenditure surveys.[8]

Differences in coverage and definitions between the price and expenditure surveys, used to establish price changes and relative weights respectively, make computation of a CLI, and therefore measurement of the difference between the CPI and a CLI, cumbersome, especially in the early years of transition. In effect, the household budget surveys used to determine

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[8] The Czech Republic is also a member of the International Monetary Fund's Dissemination Standards Bulletin Board (DSBB) which provides access to the Special Data Dissemination Standard (SDDS), the General Data Dissemination System (GDDS), and the Data Quality Reference (DQRS) sites. See http://dsbb.imf.org/

expenditure weights are conducted at a more aggregated level than the price change samples. In the example given in Table 1, prices collected at level 5 (for 738 representative items) are used to calculate the CPI, while weights from the household budget survey are only available for 192 product categories at level 4.

**Table 1.        Examples of definitions and number of strata**

| Level | CSO Code | Description | Number of strata |
|-------|----------|-------------|------------------|
| 1 | 100 000 | Food, beverages and tobacco | 10 |
| 2 | 110 000 | Food | 37 |
| 3 | 111 000 | Baked products and pastry | 103 |
| 4 | 111 300 | Bread and standard baked products | 192 |
| 5 | 111 301 | Brown bread | 738 |

It is necessary, therefore, to aggregate the price-level data to the same level as that from the expenditure survey.  Through such aggregation, we are able to link the price and expenditure surveys and, therefore, to compute the Paasche and Fisher indices for almost every year since 1989 at level 4 and above.  Because relative weights are never available at level 5 we were unable to calculate substitution within broad groups.  Neither were we able to calculate substitution across brands or outlets within narrow representative categories.  Thus, our estimates are only able to capture a portion of the substitutions made by consumers in response to shifting relative prices.[9]

The Czech Republic ( and, before 1993, Czechoslovakia)  chose an intermediate approach to price liberalization, somewhere between Poland's shock therapy and Hungary's

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[9]The same limitations apply to the Czech Statistical Office, which has no more data on expenditure patterns than was available to us from the budget surveys.  Thus, when the CSO revised CPI basket weights in 1993 and 2000 it held constant the relative shares with level-4 strata and reweighted only at level 4 and above.

gradualism. Although the Czech Republic engaged in extensive price liberalization, some prices (in particular those for electricity and housing) remain regulated and are being liberalized only gradually. Table 2 gives the dynamics of price changes and the share of controlled prices over the decade. The average annual increase in the official CPI over the period 1990 to 2000 was 9.6 per cent a year. There was, however, substantial variation in recorded inflation across various years. In January 1991, a major price liberalization was implemented and negative turnover taxes were eliminated. This move eliminated cross-subsidies that accounted for about 3 percent of GDP. As a result, prices in January 1991 were 28 percent higher than in December 1990. The effect of this price reform died out rapidly, however, and the annual inflation in 1991 as a whole was only 57 per cent. Another significant shock came from the introduction of a value added tax in January 1993 when Czechoslovakia split into the Czech and Slovak Republics. As a consequence, annual inflation increased from 11 per cent in 1992 to 21 per cent in 1993.

**Table 2.**      **Inflation Rates and Extent of Economic Reform in the Czech Republic**

| 1990 | 1991 | 1992 | 1993 | 1994 | 1995 | 1996 | 1997 | 1998 | 1999 | 2000 | 2001 |
|---|---|---|---|---|---|---|---|---|---|---|---|
| **Change in Consumer Prices (annual average)** | | | | | | | | | | | |
| 10.8 | 56.7 | 11.1 | 20.8 | 10 | 9.1 | 8.8 | 8.5 | 10.7 | 2.1 | 3.9 | 4.7 |
| **Administered Prices as Share of CPI** | | | | | | | | | | | |
| | 27.9 | 18.3 | 17.9 | 18.1 | 17.4 | 17.4 | 13.3 | 13.3 | 13.3 | 13.3 | 12.4 |
| **Private Sector as Share of GDP** | | | | | | | | | | | |
| 10 | 15 | 30 | 45 | 65 | 70 | 75 | 75 | 75 | 80 | 80 | 80 |
| **Total Trade as Share of GDP** | | | | | | | | | | | |
| 38.6 | 67.2 | 63.1 | 82.8 | 80.9 | 89.4 | 85 | 93.6 | 95.3 | 99.0 | 119.2 | 123.1 |

Source: the CSO and EBRD (2001).

For the early years of transition, the weights for the representative items comprising the consumer price index basket for the Czech Republic were determined using store turnover data. During 1993 a shift was made to weighting the basket according to results from consumer

expenditure surveys. Coordination of sample items across these surveys was poor, however, making it difficult to create an exact match between the price samples and the data used to determine weights. This matching difficulty was not resolved by the CSO until the 2001 revision of the CPI basket.

The importance of establishing proper weights can be seen in Table 3, which shows how rapidly consumption patterns shifted in Czechoslovakia in the early days of the transition.

**Table 3.      Relative Changes of Current Consumer Weights with Respect to 1989**

| Sector | 1989 | 1991 | 1992 | 1993 |
|--------|------|------|------|------|
| Foods and public catering | 40% | 36% | 34% | 34% |
| Non food products | 45% | 40% | 40% | 41% |
| Services | 15% | 24% | 26% | 25% |

Source: Czech Statistical Office.

Despite these rapid changes in consumption patterns, since the start of transition the basket used to determine changes in consumer prices in the Czech Republic has been reweighted only twice, in 1993 and 2001. Thus, there is considerable room for differences to arise between Laspeyres and Paasche indices as consumers substitute in response to changing relative prices. Table 4 presents detailed calculations of the difference between base-period and current-weight estimates of price increases as well as Fisher's exact superlative index in the Czech Republic for most years between 1991 and 1999 at the level-4 strata, one level above that at which price data was collected.[10]

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[10]Recall that we are unable to obtain current-period weights at the most detailed level (and that such weights have not shifted in official calculations. Given the rapid changes in product availability and retail distribution channels during the transition, in addition to systematic reweighting, it is also critical to update the sampling framework as outlets or referent products change. The CSO follows a linking procedure whereby new outlets and/or products are rotated into the sample as older ones become unavailable. The representative items have been defined at a fairly general level, with the individual samplers having great leeway to pick the exact items surveyed. Conversations with officials at the statistical offices lead us to believe that in the early transition years, surveyors may have frequently substituted brands

**Table 4.        Alternative Price Indices - Czech Republic** (previous year = 100)

| | Modified Laspeyres (1989 or 1993 weights) | Paasche (current weights) | Superlative (Fisher Exact) | Absolute CPI bias | Bias as % of Increase in Fisher Index |
|---|---|---|---|---|---|
| 1991 | 156.6* | 148.5 | 152.5 | 4.1 | 7.81 |
| 1992 | 111.1* | 112.1 | 111.58 | -0.48 | -4.15 |
| 1993 | 120.8* | 119.3 | 120.03 | 0.77 | 3.84 |
| 1994 | 110 | n.a**. | n.a. | n.a. | n.a. |
| 1995 | 109.1 | 108.5 | 108.8 | 0.3 | 3.41 |
| 1996 | 108.8 | 107.8 | 108.30 | 0.50 | 6.02 |
| 1997 | 108.5 | 107.8 | 108.15 | 0.35 | 4.29 |
| 1998 | 110.7 | 109.5 | 110.1 | 0.6 | 5.94 |
| 1999 | 102.1 | 101.9 | 102 | 0.1 | 5 |

Source: the CSO and authors computations.

*Uses 1989 weights.

**Because of mid-year changes of the consumer basket and weights, we are not able to calculate a Paasche index for 1994.

In the Czech Republic, where price increases have been moderate and revisions infrequent, there is a consistent pattern of the Laspeyres measures of price inflation exceeding Paasche measures.   Between 1993 and 1999, the Laspeyres index exceeded the Paasche index by an average of 9.97 per cent a year and the Fisher exact cost of living index by an average of 4.75 percent.[11]  The two first years for which we were able to calculate inflation measures present a somewhat different picture.  In 1991, which had by far the highest rate of inflation and greatest shifts in relative prices during the decade as prices were freed, the difference between

or outlets without applying formal linking procedures.

[11]Excluding 1994 when reweighting within the year made comparisons impossible.

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the Laspeyres and Paasche (or Fisher exact) indices was the greatest, with measured increases in consumer prices exceeding the increase in the cost of living by almost 8 percent. In 1992 it appears that these price increases induced supply channels increasingly to provide goods that were previously in short supply, with the result that the Paasche index actually exceeded the Laspeyres index as we suggested might happen when relative taste changes or goods availability dominates substitution caused by changing relative prices.

Over the full nine year period from 1991 to 1999, the Laspeyres Index suggested that prices rose by 236.5 per cent while the Paasche index indicated a 207.3 per cent cumulative price increase and the Fisher Exact index a 221.5 per cent increase in the cost of living.[12] These differences are substantially smaller than would be suggested by the size of relative prices changes and suggest that the observed pattern must be reflecting both demand and supply side responses that, in large part, are offsetting each other.

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[12] In fact, the difference may have been slightly larger since we were unable to calculate a Paasche Index for 1994 and, consequently, assumed that there was no bias for this year in calculating the decade average bias. The assumption of minimal bias is justified by the fact that the basket was revised during that year and should, therefore, more accurately reflect current consumption patterns.

**IV.    Computing Price and Cost-of-Living Increases at Different Strata Level**

Existing studies of substitution bias, including the results reported above, have been limited to studying substitution at what we have called level 4 or higher strata but not at the referent item level (level 5).  The inability to calculate differences at a very detailed level arises because consumer expenditure surveys typically do not provide sufficiently detailed information to recalculate weights on an annual basis.  Thus, researchers are able to capture substitution from, say, canned soup to bread, but cannot capture substitution inside a given category such as between white bread and croissants.  Reinsdorf (1996) and Bradley(1996) have used U.S. supermarket scanner data to study low-level substitution bias for a few grocery items.  These results, although limited, suggest that low-level substitution effects may also be important.

Recalling that a Laspeyres index can be computed as an aggregate of indices at lower strata levels, equation (3) can be rewritten as:

$$L_{T,T-1} = \sum_{j=1}^{n} w_{j,B}\left(P_{j,T}/P_{j,T-1}\right) = \sum_{k=1}^{n_i} w_{k,B}\left(I_{k,T}^{(l_i)}/I_{k,T-1}^{(l_i)}\right), \qquad (6)$$

where $I_{k,T}^{(l_i)}$ and $I^{(l_i)}_{\ -}$ denote Laspeyres indices computed at strata level $l_i$, and $n = n(l)$ denotes the number of representative items at strata level $l_i$.

Similarly, the aggregate Paasche index can be computed as:

$$P_{T,T-1} = \sum_{j=1} w_{j,C}\left(P_{j,T}/P_{j,T-1}\right) = \sum_{k=1}^{n_i} w_{k,C}\left(J_{k,T}^{(l_i)}/J_{k,T-1}^{(l_i)}\right), \qquad (7)$$

where $J_{k,T}^{(l_i)}$ and $J^{(l_i)}_{\ -}$ denote Paasche index computed at the $l_i$ th  strata level.

As discussed above, however, statistical offices cannot produce Laspeyres indices at more detailed strata level due to the lack of appropriate weighting information.  Thus, since

$$\sum_{k=1}^{n_i} w_{k,C}\left(J_{k,T}^{(l_i)}/J_{k,T-1}^{(l_i)}\right) \neq \sum_{k=1}^{n_i} w_{k,C}\left(I_{k,T}^{(l_i)}/I_{k,T-1}^{(l_i)}\right), \tag{8}$$

the value of Paasche indices, which must be computed using Laspeyres sub-indexes at the most detailed levels, will not be invariant to the level of aggregation at which the most detailed calculations can be done.

We have calculated Laspeyres and Paasche indices for the Czech Republic at different levels of aggregation for 1991 to 1999.  As discussed above, indexes calculated at a more aggregate strata level miss any substitution that occurred within the more detailed levels. As can be seen in Table 5, much of the variation in prices of referent items in the Czech Republic occurs at these most detailed levels.

**Table 5**.      **Relative price variation by level of aggregation (price variation of representative items at Level 5 = 1)**

| Strata Level | 1991 | 1992 | 1993 | 1994 | 1995 | 1996 | 1997 | 1998 | 1999 | 2000 |
|---|---|---|---|---|---|---|---|---|---|---|
| Level 1 | 0.28 | 0.62 | 0.18 | 0.31 | 0.27 | 0.37 | 0.39 | 0.59 | 0.36 | 0.36 |
| Level 2 | 0.46 | 0.79 | 0.43 | 0.38 | 0.45 | 0.48 | 0.52 | 0.68 | 0.56 | 0.45 |
| Level 3 | 0.64 | 0.82 | 0.65 | 0.63 | 0.56 | 0.84 | 0.72 | 0.85 | 0.69 | 0.61 |
| Level 4 | 0.74 | 0.85 | 0.82 | 0.9 | 0.72 | 0.9 | 0.82 | 0.89 | 0.74 | 0.71 |

Source: Authors' computations using data from the Czech Statistical Office.

Thus, the more aggregate the level at which calculations are performed, the closer the Paasche index should be to the corresponding Laspeyres index. This is easily seen in Table 6, which presents a Laspeyres index (invariant to aggregation level) and Paasche indices at various aggregation levels for the Czech Republic for the 1990s. There is no apparent effect of shifting relative prices at strata 1, while most of the captured effect occurs at level 2. In effect, consumers respond greatly to changes in relative prices within categories such as food or recreational goods but very little across these broad categories.[13]

**Table 6.        Paasche Indices at Various Aggregation Levels**

|  | Laspeyres | Paasche at Strata 1 | Pasche at Strata 2 | Paasche at Strata 3 | Paasche at Strata 4 |
|---|---|---|---|---|---|
| 1991 | 156.6 | 155.1 | 153.2 | 149.4 | 148.5 |
| 1992 | 111.1 | 111.6 | 112 | 111.1 | 112.1 |
| 1993 | 120.8 | 121 | 120.6 | 120.8 | 119.6 |
| 1995 | 109.1 | 109.1 | 109.1 | 108.9 | 108.5 |
| 1996 | 108.8 | 108.6 | 108.1 | 107.9 | 107.8 |
| 1997 | 108.5 | 108.4 | 108 | 107.9 | 107.8 |
| 1998 | 110.7 | 110.6 | 110.1 | 109.8 | 109.5 |
| 1999 | 102.1 | 102.2 | 101.8 | 101.8 | 101.9 |
| Mean Difference from Laspeyres |  | -0.3% | 5.5% | 7.7% | 8.5% |

Source: Authors' computations using data from the Czech Statistical Office.

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[13]This may also explain why Podpiera (2002) found no significant income bias in Czech CPI measures since he was unable to examine lower level strata.

## V. Conclusions

Although there are theoretical reasons why in a transition environment current-period weighted prices indices (Paasche indices) could exceed base-period weighted ones (Laspeyres indices), we have shown that this occurred in only one year in the Czech Republic during its transition from communism. This year was 1992, when supply distortions inherent in the shortage economy under central planning were being unwound.

Otherwise, we find the conventional pattern whereby Paasche indices fall below Laspeyres indices and, therefore, conventional measures of consumer price inflation overstate the degree of increase in true cost-of-living measures. In the Czech Republic during transition, however, the degree of this overstatement was apparently minor, averaging about 5 percent a year at the level of detail where such overstatement is conventionally measured. This degree of overstatement corresponds to those found in more stable economies such as the United States even though the apparent opportunities for substitution due to changes in relative prices were much greater in the rapidly changing transition environment. This finding suggests that, at least in part, substitution effects were offset by effects due to supply-side phenomena in which increasing prices served to eliminate shortages of given products. Such a possibility has serious repercussions for inflation management. If elimination of shortages results in inflation during the same years that the CPI understates cost of living increases (i.e. the Paasche index exceeds the Laspeyres index), central banks will tend to adopt an overly accommodating monetary policy, exacerbating the inflationary problem. Conversely, in more normal years when the bias is in the conventional direction, consumer responses to relative prices changes imply that

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responding to measured CPI changes may lead central banks to impose an overly tight monetary

policy, thereby delaying restructuring and slowing economic growth.

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