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Konstantin Gluschenko

Biases in cross-space comparisons through cross-time price indexes:

The case of Russia

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BOFIT Discussion Papers
Editor-in-Chief Iikka Korhonen

BOFIT Discussion Papers 9/2006
18.8.2006

Konstantin Gluschenko: Biases in cross-space comparisons through cross-time price indexes: The Case of Russia

ISBN 952-462-826-0
ISSN 1456-4564
(print)

ISBN 952-462-827-9
ISSN 1456-5889
(online)

Multiprint Oy
Helsinki 2006

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Konstantin Gluschenko*

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Abstract

Lacking data on price levels across locations (countries, national regions, etc.) for cross-space comparisons, researchers resort to local consumer price indexes (CPIs) over time to evaluate these levels. This approach unfortunately fails to specify, even generally, the exactness of such proxies. Worse, the method is silent on whether the results are consistent, at least qualitatively, with those obtained using actual price levels. This paper aims to find an answer empirically, using data across Russian regions. Through comparison of CPI-proxied price levels with direct evaluations of regional price levels (i.e. Surinov spatial price indexes and the costs of a purchasing power basket), biases that distort the qualitative pattern of inter-regional differences are identified. Cross-region distributions for real income (calculated with CPI-proxied and directly evaluated price levels) for several points in time are estimated and compared. The CPI-induced biases are found to generally overstate inter-regional disparities.

JEL Classification: C43**,** E31, P22, R19

Keywords: consumer price index, spatial price index, real income, nonhomothetic preferences, Russia, Russian regions

Acknowledgements: This work was done during my stay at the Bank of Finland Institute for Economies in Transition (BOFIT) as a visiting researcher. I thank BOFIT and its staff for their hospitality and for providing an excellent research environment. My gratitude further extends to Vladimir Bessonov, Iikka Korhonen, and the participants of BOFIT seminars for their insightful and helpful discussions.

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Konstantin Gluschenko

Biases in cross-space comparisons through cross-time price indexes: The case of Russia

Tiivistelmä

Koska tutkijoilla ei aina ole käytössään tietoa eri maantieteellisten alueiden (maiden, hallinnollisten alueiden jne.) hintatasosta, he käyttävät usein paikallisia kuluttajahintaindeksejä hintatasojen vertailuun. Tämän lähestymistavan ongelma on kuitenkin, että indeksin tarkkuutta ei pystytä määrittämään edes summittaisesti. Vielä ongelmallisempaa on se, että indeksejä vertailtaessa ei yleensä pohdita, ovatko tulokset edes kvalitatiivisesti samoja kuin hintatasoja suoraan vertailtaessa. Tässä tutkimuksessa ongelmaa lähestytään empiirisesti käyttäen Venäjän eri alueita koskevia tilastotietoja. Vertaamalla kuluttajahintaindeksejä alueellisiin hintatasotietoihin (Surinovin spatiaalisiin hintaindekseihin tai ostovoimakorin hintaan) pystytään identifioimaan alueiden välisten hintatasojen vertailussa syntyvät harhat. Lisäksi tässä työssä lasketaan reaalinen tulotaso eri alueilla eri ajankohtina käyttäen sekä kuluttajahintaindeksejä että suoraan havainnoituja alueellisia hintatasoja. Alueellisten kuluttajahintaindeksien käyttö alueiden välisten tuloerojen laskemiseen näyttää liioittelevan tuloeroja.

Asiasanat: kuluttajahintaindeksi, spatiaalinen hintaindeksi, reaalinen tulotaso, ei-homoteettiset preferenssit, Venäjä, Venäjän alueet

1 Introduction

Many areas of economic study require knowledge of aggregated local price levels across countries, national regions, or other geographical entities. Country price levels can be dealt with directly in testing for purchasing power parity (PPP). Integration of domestic markets can be studied with regional or city price levels. Local price levels can also be applied in converting nominal monetary indicators (incomes, wages, consumption, etc.) into real monetary indicators for spatial comparisons, particularly analysis of spatial inequality. The convergence hypothesis in the context of economic growth should be tested in terms of spatially comparable real values.

Price comparisons differ qualitatively across space and across time (see e.g. ILO, 2004, p. 495). Though not nearly as voluminous as the literature on temporal index numbers, the literature on spatial price indexes is also extensive.[1] A number of papers use spatio-temporal index numbers in discussion of local price levels.[2]

However, data on local price levels are in fact quite rare. Lacking local price level data, researchers have traditionally dealt with this problem in one of two ways. The first is to use nominal values. Barro and Sala-i-Martin (2004), for example, test for (income) convergence across the US states, Japanese prefectures, and European regions. Their data descriptions suggest that the incomes used are, in essence, nominal.[3] Thus, given the abundant evidence of wide variability of consumer prices across US cities (and impliedly states), one cannot confidently conclude from the results whether a convergence of welfare actually occurred. The second popular way is to evaluate local price levels at a given point in time through changes in the levels from a base point in time (provided that the price levels at the base point are known). That is, local price levels are approximated by local con-

[1] Diewert (1999) provides a review with numerous references.
[2] Hill (2004) offers a rather good summary.
[3] In the context of both temporal and spatial comparisons, the term “real” is somewhat confusing. “We have computed real income by dividing the nominal figures on personal income [by US state] by the national values of the consumer price index,” write Barro and Sala-i-Martin (2004), p. 497. In one sense, transformed incomes can be called real in that they are uniformly deflated. (Equally, or even better, this can be achieved with dividing the incomes in states by the national average.) Whatever the case, they are not real in the sense of spatial comparability as they are measured in monetary units that have different purchasing power across states. For example, per capita income in 2002 was 104.9% of the US average in Alaska and 94.6% in Nebraska (calculated from www.bea.gov/bea/regional/spi). Is it in fact correct to say that in 2002 real income in Alaska was higher than in Nebraska when the cost of living, according to ACCRA (2002), was 120.5% to 131.3% of the national average across Alaskan cities and 92.2% to 96.4% across cities of Nebraska?

sumer price indexes (CPIs). This method is found in most of papers on testing PPP,[4] as well as studies of domestic market integration (e.g. Cechetti, Mark and Sonora, 2002). In the absence of data on initial price levels, PPP is either assumed to hold on average in the base period (i.e. relative initial price levels are assumed to equal roughly 1), or, more commonly, analyzed in its relative form, comparing changes in price levels across locations rather than levels themselves. The use of local CPI as a measure of local cost of living is also commonly applied in comparative studies to deflate nominal local indicators, so converting them into spatially comparable real values. This is especially attractive in the case of the post-socialist countries, as prices across regions of a country before transition can readily be deemed equal due to centralized pricing. In the case of Russia, this approach has been adopted by e.g. Dolinskaya (2002), Kwon and Spilimbergo (2005), and Solanko (2003).

Knowing the value of a local price level at some point in time would seem to suggest that we could extend it to any other point in time, multiplying the initial level by rise in prices in the location since that time and then compare the obtained price levels across locations. Unfortunately, this notion is fatally flawed because it ignores the conceptual inconsistency between price levels involved in comparisons across space and across time. Even where researchers acknowledge this inconsistency, they still use this method for an approximation as they have no other way to estimate local monetary indicators in spatially comparable real terms. As a result, they cannot specify the exactness of their approximations. Worse, the method is silent on whether the results are consistent, at least qualitatively, with those that would be obtained with the use of actual price levels.

This paper aims to answer these questions empirically, using data across Russian regions. Theoretically, a bias yielded by the CPI approximation is a deviation of CPI-proxied regional price level from a "perfect" estimate. In such case, the best way to obtain the benchmark would be to use a method that provides as close as possible approximation of this "perfect" estimate.

There are two impediments to this approach, however. While the Divisia index can be deemed as "perfect" for comparisons across time, there is nothing similar for the case of cross-space comparisons. Thus, arbitrariness is inevitable in the choice of a "best" benchmark. The second, and far more practical, problem is that of data availability. Prices across Russian regions are not released, although conceivably one could purchase such informa-

[4] Rogoff (1996) and Taylor and Taylor (2004) surveyed some of the more important works.

tion directly from Russia's statistical agency, Rosstat (formerly Goskomstat). The price weights involved in computation of Russian regional CPIs, however, would not be provided as Rosstat treats them as in-house information.[5] Thus, even if we chose one of these methods to obtain the "best" benchmark, it would still be beyond our abilities to calculate the benchmark.

To overcome this, we estimate biases in CPI-proxied regional price levels as the deviations from available direct evaluations of regional price levels. These are spatial price indexes computed by Surinov (1999) for January 1997 and January 1998 with the same data used in computation of Russian CPIs and the cost of a fixed basket of (83) goods and services for inter-regional comparison of population's purchasing capacity published monthly by Goskomstat/Rosstat since January 2002. The cost of a staples basket (a component of the subsistence minimum) is used to verify the results.

CPI-proxied price levels are found to be biased sufficiently to distort the qualitative pattern of inter-regional differences. When cross-region distributions of different versions of real income (estimated with the use of the CPI-proxied and directly evaluated price levels) are compared, the CPI-induced distortions of the distribution shape and statistics become evident. Although CPI-induced biases have different directions across regions, they generally overstate inter-regional disparities. Numerical experiments suggest that even if a "perfect" CPI (the Divisia index) is used and favorable conditions other than homotheticity of consumer preferences are provided, biases in the CPI-proxied price levels are sufficient.

The implications of such findings extend far beyond Russia. The Russian CPI is conceptually similar to the CPI in most countries, including EU member countries. The Harmonized Index of Consumer Prices (HICP) for EU countries covers the same set of goods and services across countries, but with country-specific weights. Thus, national HICPs can be seen as comparable with Russian regional CPIs, while the European Index of Consumer Prices (EICP) is similar to the Russian national CPI. The effects found are inherent in spatial comparisons that use local CPIs within the EU, and even more so in international comparisons, where the CPI baskets vary across countries. Such effects can be hard to detect during periods of low inflation and may only express themselves in the medium or long run.[6] In this sense, Russia is interesting because it provides a kind of acceler-

[5] De Masi and Koen (1995) are apparently the only authors who succeeded in obtaining such data. Their working paper reports weights in the Russian national CPI for 1993.
[6] On the other hand, given small changes in local prices, there is no need in CPIs to correct inter-location figures.

ated natural experiment with economic time compressed by transition. Effects normally evident only in the long run in established market economies are readily apparent during transition.[7]

The remainder of the paper is organized as follows. In the next section, the methodology and data are described. Section 3 reports results on comparison of the CPI-proxied price levels with directly estimated regional price levels. Examples of applying different measures of the regional costs of living to estimation of real incomes are provided in section 4. Section 5 offers the results of numerical experiments. Section 6 discusses the main reasons for biases in CPI-proxied price levels. Conclusions and policy implications are handled in section 7.

2 Methodology and data

2.1 Preliminaries

The following terminology is used. A *spatial price index* (SPI), P_{rst}, means a price level in location r relative to location s (reference location) at time t. The term *direct spatial price index* (DSPI) is used in reference to an SPI calculated for t with the use of local prices for the same point in time. The term *indirect spatial price index* (ISPI) applies to an SPI extrapolated from a direct SPI for a time point t_0 (the base DSPI) to point t with the use of local CPIs characterizing price changes over the time interval from t_0 to t ("CPI-proxied" SPI).

For our purposes, s is fixed; Russia as a whole is taken as the reference location, denoted as $s = 0$. Thus, SPIs are constructed with the "star system," using the whole of the country as the star location (an artificial location). Biases in CPI-proxied SPIs are estimated as deviations of ISPIs from DSPIs.

Due to inadequate data available for the Chukot Autonomous Okrug, Chechen Republic, and Republic of Ingushetia, we only consider 86 of Russia's 89 regions. In addi-

[7] For example, the price level in Sweden increased 65-fold since 1831 through 2004, that is, over 173 years (www.scb.se/templates/tableOrChart____33838.asp). In Russia, it took only 12 years, since December 1993 through December 2005, for the price level to increase 63-fold (not to mention 1992-1993, when the figure was 246).

tion, nine autonomous *okrugs* are part of other regions (*oblast* or *krai*). In these cases, we only use the data for the entire *oblast* or *krai*. This gives us 77 spatial observations. For ease of presentation, the regions are aggregated to eleven "macroregions," economic areas (*economicheskiy rayon*),[8] following the traditional economic-geographic zoning of Russia. We incorporate the Kaliningrad Oblast, usually treated as a separate area, into the North-Western area.

To aggregate some regional indexes Z_l, the share of the population of economic area r, $N_{l(r)}/N_r$, is used a weight of a region; $l(r)$ indexes regions belonging to area r:

$$Z_r = \sum_{l(r)} \frac{N_{l(r)}}{N_r} Z_{l(r)} . \tag{1}$$

Aggregating over all economic areas in the same manner is expected to yield the national index Z_0, which we apply as our reference value.

2.2 Indirect spatial price index

Absolute price level P_{rt_0} in location r is assumed to be known for time point t_0, as well as the change in the price level over the time interval from t_0 to t, $I_r(t_0, t)$. From this we compute the absolute price level for time point t as $P_{rt} = P_{rt_0} \cdot I_r(t_0, t)$. Now that we have relevant data regarding location s, the indirect spatial price index at time t for locations r and s is given by

$$P_{rst} = \frac{P_{rt_0} \cdot I_r(t_0, t)}{P_{st_0} \cdot I_s(t_0, t)} = P_{rst_0} \frac{I_r(t_0, t)}{I_s(t_0, t)} . \tag{2}$$

Thus, P_{rst_0} is the DSPI for t_0.

Next, the change in price level is measured by a CPI calculated by the chain method such that

$$I_r(t_0, t) = \prod_{\tau = t_0 + 1}^{t} I_r(\tau - 1, \tau) . \tag{3}$$

[8] The English version of this term and the English names of the areas are drawn from Goskomstat (1998), p. 56. This source also provides descriptions of the economic areas.

Briefly, the Russian CPI is constructed as follows.[9] The regional CPIs are produced at a monthly frequency. The price index number formula used is

$$I_l(\tau-1,\tau)=\sum_k w_{kl\theta}\frac{p_{kl\tau}}{p_{kl,\tau-1}}, \tag{4}$$

where l indexes regions of the country, τ indexes months, k indexes commodities, θ is a weight reference period, $w_{kl\theta}$ is an expenditure share, and $p_{kl\tau}$ is a commodity price. The weights ($w_{kl\theta}$) are derived from the previous year (θ) household budget survey. The survey is conducted annually; and the weights are updated every year. The index of the form (4) is known as the Young index (with monthly prices and annual base year weights).[10]

In addition to the overall CPI, Goskomstat/Rosstat publishes three its sub-categories: foods (including beverages and alcohol), manufactured goods, and services. The coverage of commodities varies over time. For example, the Russian CPI for 1996 covered 288 items: 73 foods, 144 industrial goods, and 71 services; in 2002, the total number of items was 411, including 100 foods, 214 industrial goods, and 97 services.

The national CPI is constructed in the same way, using national expenditure shares and national elementary price indexes. The latter is a weighted average of the regional elementary indices, the weights being the regional shares of Russia's population, N_l/N:

$$I'_0(\tau-1,\tau)=\sum_k w_{k0\theta}\sum_l \frac{N_l}{N}\cdot\frac{p_{kl\tau}}{p_{kl,\tau-1}}. \tag{5}$$

Comparing this formula with (1), we see that they exploit different ways of aggregation over regions. To provide consistency with economic area CPIs obtained by Formula (1) with the use of $I_l(\tau-1,\tau)$ as Z_l, the reference value computed in the manner of (1) is used instead of the official values yielded by (5). Thus,

$$I_0(\tau-1,\tau)=\sum_l \frac{N_l}{N} I_l(\tau-1,\tau) \tag{6}$$

is taken in the chain (3) for Russia as a whole. While the values of $I_0(\tau-1,\tau)$ differ from those for $I'_0(\tau-1,\tau)$, the differences are typically less than one percent.

Before the price liberalization at the beginning of January 1992, consumer prices were fixed and almost uniform throughout Russia due to the centralized pricing. Taking

[9] See Goskomstat (1996) and Bessonov (1998) for details.

[10] To underscore the fact that the weight reference period is prior to the comparison period, τ, Goskomstat/Rosstat refers to this index as a "modified Laspeyres index," while Eurostat calls it a "Laspeyres-type index."

December 1991 as the base period (t_0 = 1991:12), the base DSPI in Formula (2), P_{rst_0}, approximately equal 1 for each r. Thus, the ISPI looks like

$$P_{r0t}^{(\pi)} = P_{r0,1991:12} \frac{I_r(1991:12,t)}{I_0(1991:12,t)} \equiv P_{r0,1991:12} \cdot I_{r0}(1991:12,t) = I_{r0}(1991:12,t).^{11} \qquad (7)$$

We offer this formula with two caveats. The first is that administratively set prices varied across regions due to "zone prices" set for certain foods. The three administrative zones exhibited modest differences in prices. Neglecting these differences results in minor understatement of the base DSPIs for the more remote regions, especially Siberia and the Far East.

Second, market pricing was already starting to happen at the close of the planned economy era. Thus, there may be differences in consumer prices across regions by the end of 1991. While a better base period might then be December 1990, there are unfortunately no official CPIs by region for 1991 (and the unofficial CPIs are quite unreliable). We simply argue here that market pricing still had a minor effect on prices *involved in the CPI* in 1991. Gluschenko (2000) discusses this issue, but for our purposes here it suffices to note that the prices surveyed in 1991 came from state-run stores. Such stores were an overwhelming majority of outlets at that time; they sold, for the most part, goods with fixed prices. Goods with market-set prices were sold at that time in city market squares, kiosks, through street trade, etc. Moreover, most services covered by the CPI (e.g. housing and utilities) had fixed prices. Nonetheless, doubts about the uniformity of prices across regions in the end of 1991 are not unreasonable. What we can say is that the cross-region divergence of prices in the base period was likely quite modest.

To compute the ISPIs, the official monthly CPIs (overall CPI, CPI-food, CPI-manufactured goods, and CPI-services) by region provide the raw data. Data for January 1992 through December 1995 are obtained directly from Goskomstat; data from January 1996 on are drawn from Goskomstat/Rosstat's monthly bulletins *Socio-Economic Situation of Russia*.

[11] Modifications of this ISPI with other base periods and base DSPIs are used in the following discussion.

2.3 Direct spatial price indexes

The Surinov DSPI. Surinov (1999) calculated DSPIs for January 1997 and January 1998 by Russian region, applying the star system with Moscow as the star region. His index number formula is a Young index:

$$P_{lst}^{(S)} = \sum_k w_{k0\theta} \frac{p_{klt}}{p_{kst}}, \tag{8}$$

where the weights are the same as those used in the Russian national CPI (5) for respective periods (coverage of goods and services coincides with the Russian CPI); s is fixed and corresponds to Moscow. As with the CPI, the overall DSPI and its three sub-categories are reported.

In addition to the regional indexes, Surinov (1999) reports DSPIs for Russia as a whole relative to Moscow, $P_{0st}^{(S)}$. The latter is constructed in much the same way as (8) with national prices p_{k0t} instead of p_{klt}. Here, the national prices are weighted averages of regional prices with weights N_l/N. Since the commodity weights are uniform across regions, this is equivalent to aggregating regional indexes (8) in the manner of (1).

For convenience, after aggregating the regional indexes into the economic area indexes by (1), the indexes relative to Moscow are converted into indexes relative to Russia as a whole, applying renormalization $P_{r0t}^{(S)} = P_{rst}^{(S)} / P_{0st}^{(S)}$. We refer to the DSPI of this form as the Surinov DSPI. As is known, the Young index is not transitive (i.e., $P_{rlt} \cdot P_{lst} \neq P_{rst}$). Hence, the above transformation diminishes the accuracy of the DSPI.[12] However, experimental comparisons of ISPIs relative to Moscow with original DSPIs (8) yielded the same results as comparisons of the indexes relative to Russia as a whole. Thus, discrepancies between approximate and exact values of the DSPIs relative to Russia as a whole are within the accuracy of the original Surinov indexes.

Basket DSPIs. In January 2002, Goskomstat has started publishing the cost of a fixed basket of goods and services for inter-regional comparison of population’s purchasing capacity by region and for the entire country (“purchasing power basket”). This basket covers 83 commodities, including 30 foods, 41 manufactured goods, and 12 services. It has uniform quantities of commodities, q_k, across regions and time. The cost of the basket (ag-

[12] The exact formula would be (8) with p_{kst} substituted for p_{k0t}.

gregated by economic area), $P_{rt}^{(\mathrm{PP})} = \sum_k q_k p_{krt}$, is in fact an absolute price level in r, and the DSPI is constructed as

$$P_{r0t}^{(\mathrm{PP})} = P_{rt}^{(\mathrm{PP})} / P_{0t}^{(\mathrm{PP})} . \tag{9}$$

We referred to this as the PP basket DSPI or the relative cost of PP basket.

There is another basket used by the Russian official statistics, a staples basket. In the period 1992-1996, it included 19 foods. Between January 1997 and May 2000 it contained 25 foods. Starting in June 2000, it incorporates 33 foods. The costs of these baskets are calculated in much the same way as above with relevant commodity sets $\{k\}$ and quantities q_k. We refer to the relative costs, obtained similarly to (9), of the basket of 25 and 33 staples as “the cost of basket-25” ($P_{r0t}^{(25)}$) and “the cost of basket-33” ($P_{r0t}^{(33)}$), respectively.

Raw data on the costs of the purchasing power basket and basket-33 are drawn from Goskomstat/Rosstat’s monthly bulletins *Socio-Economic Situation of Russia*. The costs of the basket-25, including those calculated retrospectively for February 1992 through December 1996, were obtained directly from Goskomstat.[13] The basket-25 represents about one third of the food items in the CPI and, as calculated from data reported by de Masi and Koen (1995), 56.7% of the 1993 weights in the national CPI-food.

Yemtsov (2005) uses official data on a subsistence minimum to convert nominal regional incomes across Russian regions into real terms. Indeed, the subsistence minimum would be an interesting proxy of the regional costs of living if its basket were region-specific. Unfortunately, before 2000 the value of the subsistence minimum was the cost of the staples basket plus other expenses expressed as its percentage, which was uniform across regions. Thus, the relative subsistence minimum coincides with the relative cost of the staples basket. Since 2000, a new methodology has been used. The basket now includes specific manufactured goods and services rather than the crude estimate of the total expenses for them. Even so, this basket is still uniform across regions.

[13] See Gluschenko (2003) for a description of this data set.

3 Empirical results

3.1 Indirect SPI vs. Surinov SPI

Table 1 reports ISPI (7) for 1997:1 and 1998:1 and the Surinov DSPI. For foods, the relative cost of basket-25 is reported as well. Since Surinov (1999) reports his indexes as integer percents, the data are displayed to the second decimal place.

The table shows sizable discrepancies between the ISPI and DSPI. The ISPI overstates inter-spatial gaps when it suggests the overall price level varies across economic areas of Russia from 81% of the national average to 153%. The DSPI, in contrast, narrows this range to 92-136%. The pattern is similar for sub-categories. The ratio of the highest price level to the lowest one is 2.4 as estimated by the ISPI and 1.7 by the DSPI for foods (1.8 for the staples basket); 1.8 and 1.4, respectively, for manufactured goods; and 2.9 and 1.8 for services (maximum ranges are taken for 1997 and 1998).

Table 1. Comparison of indirect SPI with Surinov direct SPI

Economic area	Overall		Foods			Manufactured goods		Services	
	ISPI	DSPI	ISPI	DSPI	Staples*	ISPI	DSPI	ISPI	DSPI
1997:1									
Northern	0.95	1.08	0.81	1.11	1.15	1.03	0.98	0.86	1.21
North-Western	0.81	0.99	0.89	0.94	0.93	0.80	1.00	0.61	1.09
Central	0.95	0.94	1.00	0.99	0.97	0.90	0.96	0.76	0.73
Volga-Vyatka	0.82	0.97	0.84	0.93	0.87	0.89	0.92	0.85	1.17
Central Black Soil	0.98	0.92	0.95	0.88	0.89	1.20	0.91	0.72	1.01
Volga	0.86	0.95	0.96	0.93	0.89	0.87	0.95	0.77	0.99
Northern-Caucasus	0.91	0.95	0.67	0.90	0.91	1.01	1.01	0.83	0.91
Urals	1.00	1.03	1.07	0.98	0.96	0.85	1.00	1.05	1.22
Western Siberian	1.25	1.05	0.98	1.06	1.02	1.46	1.05	1.45	1.02
Eastern Siberian	1.17	1.08	1.63	1.14	1.20	0.89	1.03	1.18	0.96
Far Eastern	1.53	1.36	1.47	1.47	1.57	1.45	1.25	1.78	1.26
1998:1									
Northern	0.93	1.04	0.79	1.06	1.11	1.01	0.98	0.79	1.14
North-Western	0.81	0.99	0.90	0.95	0.97	0.80	0.98	0.58	1.16
Central	0.95	0.94	1.00	1.00	1.02	0.90	0.96	0.75	0.73
Volga-Vyatka	0.81	0.95	0.82	0.90	0.85	0.89	0.94	0.84	1.14
Central Black Soil	0.99	0.91	0.97	0.89	0.89	1.20	0.89	0.74	1.06
Volga	0.88	0.95	0.95	0.92	0.89	0.87	0.96	0.86	1.05
Northern-Caucasus	0.94	0.95	0.69	0.92	0.94	1.02	0.99	0.89	0.92
Urals	0.98	1.02	1.06	0.97	0.93	0.84	1.00	1.00	1.22
Western Siberian	1.24	1.03	0.98	1.05	1.01	1.46	1.05	1.38	0.92
Eastern Siberian	1.14	1.06	1.61	1.12	1.13	0.87	1.04	1.13	0.92
Far Eastern	1.53	1.33	1.43	1.40	1.55	1.47	1.24	1.79	1.34

* The relative cost of basket-25, $P^{(25)}_{r0t}$.

While the DSPI values look reasonable, some values of ISPI are immediately suspect, such as those evidencing that the food price level in Eastern Siberia was 60% higher than the national average (and exceeded the price level in the Far East which includes a number of difficult-to-access regions such as Yakutia, Kamchatka, and the Magadan Oblast), or that the Western Siberian level of prices for manufactured goods was half as much again the national level. These figures directly conflict with observations of actual prices at that time.

Even the qualitative pattern yielded by the ISPI is dissimilar to that yielded by the DSPI. The rankings of economic areas by one index or another turn out quite different. By contrast, the relative cost of basket-25 is much more consistent with the DSPI for foods, despite much smaller coverage. The area ranks by these two indexes differ by no more than one, except for the Volga-Vyatka economic area in 1997 and the Urals in 1998, where the difference equals two. Quantitatively, the divergence of the relative cost of basket-25 from the DSPI ranges from –6.5% to 5.3%, excluding the Far East, where the divergence is 6.8% in 1997 and 10.7% in 1998.

Table 2 gives the percent values of ISPI biases computed as $(P_{r0t}^{(\pi)} - P_{r0t}^{(S)}) / P_{r0t}^{(S)}$. When we take the 5% band as a conventional "confidence interval," about a quarter of biases fall within this band. There are three such cases (in each period) for the overall ISPI, four cases for the ISPI-food, three cases for the ISPI-manufactured goods, and two cases (with one more in 1998) for the ISPI-services. There is no single economic area where all four ISPIs might be deemed unbiased. The Central economic area comes closest in this respect.

Table 2. ISPI bias percentages, (ISPI – DSPI)/DSPI

Economic area	**Overall ISPI**		**Foods**		**Manufactured goods**		**Services**	
	1997:1	**1998:1**	**1997:1**	**1998:1**	**1997:1**	**1998:1**	**1997:1**	**1998:1**
Northern	-11.6	-11.3	-27.3	-26.0	4.6	3.0	-29.3	-31.0
North-Western	-17.8	-17.9	-4.5	-4.8	-20.5	-18.1	-43.5	-49.8
Central	0.2	0.1	0.9	-0.3	-6.1	-6.9	4.4	3.2
Volga-Vyatka	-15.8	-15.5	-9.6	-8.8	-3.7	-5.0	-27.1	-26.5
Central Black Soil	6.7	8.4	8.1	9.0	31.9	34.3	-28.3	-29.8
Volga	-9.2	-8.2	3.6	3.2	-8.3	-7.7	-22.2	-17.5
Northern-Caucasus	-4.2	-1.4	-26.0	-25.6	-0.6	2.5	-8.6	-3.1
Urals	-2.9	-3.4	8.4	8.5	-14.6	-16.2	-14.4	-18.2
Western Siberian	18.6	20.9	-7.0	-6.1	39.4	39.2	43.0	50.3
Eastern Siberian	8.1	8.0	43.2	44.0	-13.9	-16.3	22.5	22.9
Far Eastern	12.9	15.3	-0.5	2.0	16.1	18.4	41.1	33.3

Figure 1 provides a more detailed pattern of the biases, reporting their empirical distributions across regions. The figure contains a chart for each kind of the ISPI: overall, foods, manufactured goods, and services. Each chart plots kernel estimates of the density of biases for 1997:1 and 1998:1. Table 3 tabulates summary statistics of these distributions.

Figure 1. Cross-region empirical distributions of ISPI biases

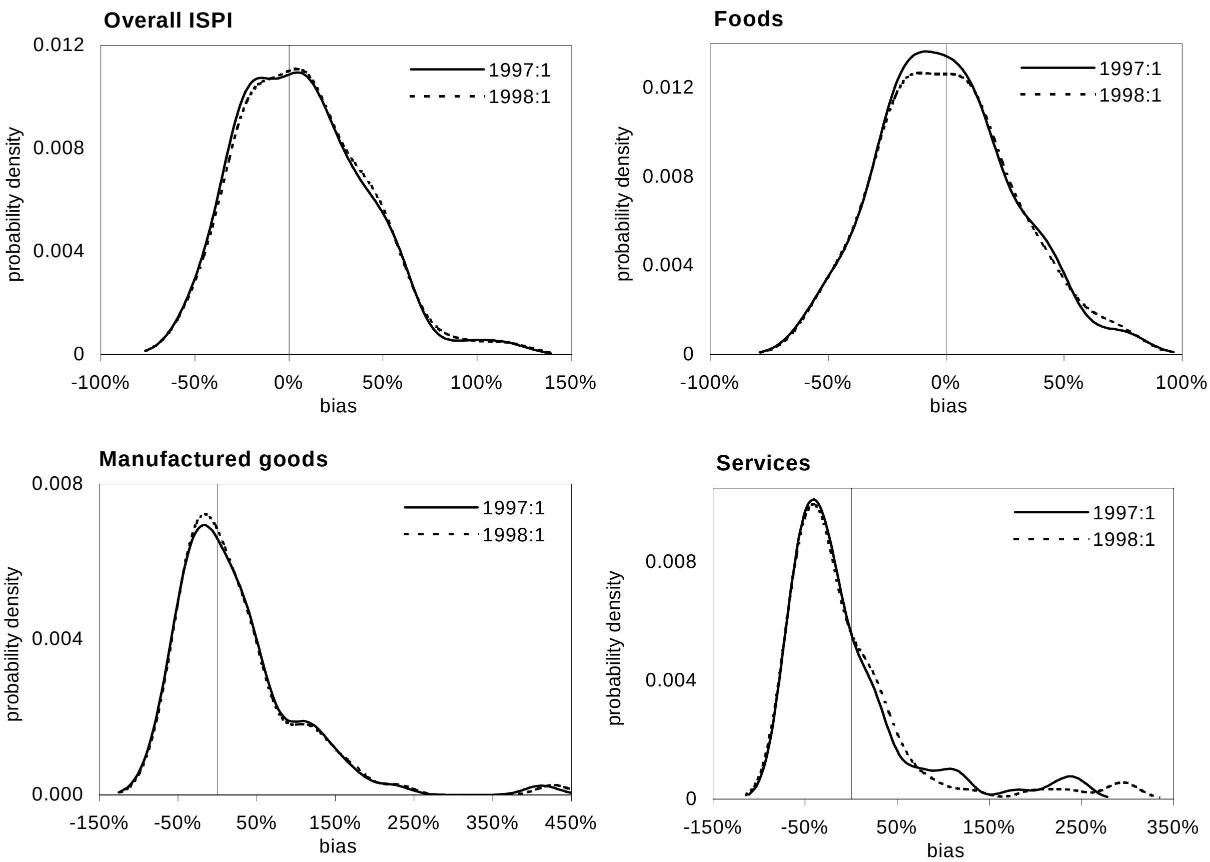

Table 3. Summary statistics for cross-region bias distributions (%)

Kind of ISPI	Mean	Median	Standard deviation	Minimum	Maximum
		1997:1			
Overall	6.8	4.8	33.1	-51.1	114.4
Foods	0.1	-1.9	28.3	-58.2	75.7
Manufactured goods	23.2	1.5	78.0	-82.4	412.1
Services	-5.4	-26.2	69.4	-77.6	244.6
		1998:1			
Overall	7.9	5.9	33.2	-51.4	116.9
Foods	0.7	-0.6	28.8	-57.5	73.6
Manufactured goods	23.8	4.6	79.6	-82.7	427.3
Services	-5.0	-28.8	74.6	-78.5	298.7

The bias of the overall ISPI ranges across economic areas from –18% to 21%, the range for foods is –27% to 44%, the range for manufactured goods is –20% to 39%, and the range for services is –50% to 50%. Thus, the discrepancies between the indirect and direct estimates of price levels run as high as 1.5 times on the level of economic areas. On the regional level, however, the bias ranges are much wider. The widest range is that for manufactured goods: –83% to 427%, a five-fold variation. This results in a severe distortion of cross-space comparisons. As seen from Tables 2 and 3, the distortion changes in time: some biases rise in absolute value while others diminish.

As mentioned in Section 2.2, there is no certainty that the base price levels involved in calculation of ISPIs are uniform across regions or economic areas. Thus, the question arises: Is it an inadequacy of the base DSPIs that is responsible for the biases? Hypothesizing that this is the case, i.e. $P_{r0t}^{(\pi)} = P_{r0t}^{(S)}$, from (7) we have $P_{r0,1991:12} = P_{r0t}^{(S)} / I_{r0}(1991:12, t)$, an estimate of possible base DSPI. Table 4 reports these estimates for t = 1997:1.

Table 4. Possible base DSPIs, DSPI/ISPI

Economic area	Overall	Foods	Manufactured goods	Services
Northern	1.13	1.38	0.96	1.42
North-Western	1.22	1.05	1.26	1.77
Central	1.00	0.99	1.06	0.96
Volga-Vyatka	1.19	1.11	1.04	1.37
Central Black Soil	0.94	0.92	0.76	1.39
Volga	1.10	0.97	1.09	1.29
Northern-Caucasus	1.04	1.35	1.01	1.09
Urals	1.03	0.92	1.17	1.17
Western Siberian	0.84	1.08	0.72	0.70
Eastern Siberian	0.92	0.70	1.16	0.82
Far Eastern	0.89	1.00	0.86	0.71

If the hypothesis were true, the estimates of base DSPI would differ in a range of no more than a few percentage points (see Section 2.2). Yet the pattern suggested by Table 4 looks highly improbable.

First, there was no such wide spatial variation of prices for goods in the end of 1991, namely, from 70% to 138% of the Russian average for foods, or from 72% to 126% for manufactured goods. The range of 70% to 176% for services seems just as unlikely. Nevertheless, it is impossible to say for sure, because fixed prices for services were very low in 1991, which could result in small weights for them. Lacking data on the CPI weights, there is no way to clarify the issue.

Second, the spatial distribution of the relative price levels has nothing common with reality of that time. A couple cases suffice. If market prices played a significant role in the end of 1991, levels of prices for goods would be the highest in the Far Eastern economic area due to high shipping costs as well as the zone pricing (recall that this area includes a number of remote northern regions). But the figures in Table 4 suggest that prices for foods were about the national level there. Moreover, prices for manufactured goods turned out to be below the national level. Another example is Northern Caucasus, where foods should be cheaper than almost elsewhere in Russia. In fact, the estimated “base DSPI” is a one third higher than that of Russia as a whole. Thus, it can be concluded that inaccuracy of the base DSPIs cannot be responsible for the biases found, or at minimum, its role is not fundamental.

The use of two benchmarks (DSPI for 1997:1 and 1998:1) offers yet another way to identify the existence of biases. The ISPI is adjusted for the bias accumulated by 1997:1 through the use of DSPI for 1997:1 as the base DSPI. Once adjusted, the ISPI for 1998:1 is compared with the respective DSPI. A difference between them is seen to characterize the bias over the 12-month time span. Thus, the adjusted ISPI is $P'^{(\pi)}_{r0t} = P^{(S)}_{r0,1997:1} \cdot I_{r0}(1997:2, t)$; and t = 1998:1. Table 5 presents the results; for ease of comparison, DSPIs for 1998:1 from Table 1 are reported as well.

Table 5. ISPIs adjusted for 1997:1 bias vs. Surinov DSPIs for 1998:1

Economic area	**Overall**		**Foods**		**Manufactured goods**		**Services**	
	Adjusted ISPI	**DSPI**	**Adjusted ISPI**	**DSPI**	**Adjusted ISPI**	**DSPI**	**Adjusted ISPI**	**DSPI**
Northern	1.05	1.04	1.08	1.06	0.96	0.98	1.12	1.14
North-Western	0.99	0.99	0.95	0.95	1.01	0.98	1.03	1.16
Central	0.94	0.94	0.99	1.00	0.95	0.96	0.72	0.73
Volga-Vyatka	0.96	0.95	0.91	0.90	0.93	0.94	1.16	1.14
Central Black Soil	0.93	0.91	0.90	0.89	0.91	0.89	1.03	1.06
Volga	0.97	0.95	0.92	0.92	0.97	0.96	1.11	1.05
Northern-Caucasus	0.98	0.95	0.93	0.92	1.02	0.99	0.97	0.92
Urals	1.01	1.02	0.98	0.97	0.98	1.00	1.17	1.22
Western Siberian	1.05	1.03	1.06	1.05	1.05	1.05	0.96	0.92
Eastern Siberian	1.06	1.06	1.12	1.12	1.01	1.04	0.92	0.92
Far Eastern	1.36	1.33	1.44	1.40	1.26	1.24	1.27	1.34

Bearing in mind that the Surinov's (1999) data are accurate to the second decimal place, a disparity between the adjusted ISPI and the 1998 DSPI in the range of about one percentage point can be attributed to a rounding error. About a half of disparities (five out of eleven) fall in this range if the overall indexes are dealt with. In the case of foods, nine fall within the range. However, only two to four indices can be deemed as coinciding in a pair if the case in hand is services or manufactured goods. The remaining disparities should be assigned to biases accumulated during 12 months. The maximum percent discrepancy between overall indices is 3.2% (the Northern Caucasus area), that for foods is 2.9% (the Far Eastern area), and that for manufactured goods is 3.1% (the North-Western area). For services, the figure is almost as four times as high, 11.2% (the North-Western area).

As compared to Table 2, the biases are rather small. Obviously, the reason is that it is a short time span of low inflation by Russian standards. Inflation over 1998:1 through 1998:1 was 10.1% (varying from 7.8% to 14.1% across economic areas), with rise in prices for foods, manufactured goods, and services equaling 8.1%, 7.7%, and 20.5%, respectively. Moreover, aggregation of regions into economic areas markedly smoothes biases. Figure 2 reports empirical distributions of biases of the adjusted ISPI across regions. Table 6 contains summary statistics of these distributions.

Figure 2. Cross-region empirical distributions of biases of ISPI adjusted for 1997:1 bias; 1998:1.

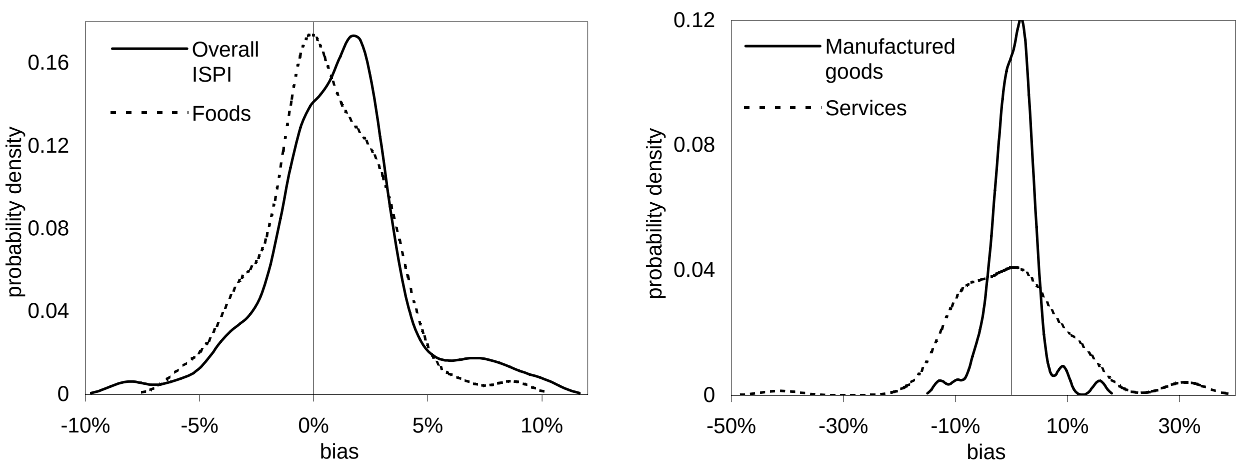

Table 6. Summary statistics for cross-region bias distributions; ISPI adjusted for 1997:1 bias; 1998:1 (%)

Kind of ISPI	Mean	Median	Standard deviation	Minimum	Maximum
Overall	1.1	1.1	3.0	-8.1	10.0
Foods	0.5	0.3	2.6	-5.9	8.7
Manufactured goods	0.4	0.6	4.0	-12.8	15.7
Services	0.1	-0.3	10.8	-41.1	32.3

Figure 2 and Table 6 suggest that regional biases of the adjusted ISPI can be much greater than in the case of economic areas. The bias of the overall ISPI ranges from –8% to 10% across regions; the range of the ISPI for services is –41% to 32%. Dealing with seven particular regions constituting Western Siberia, Gluschenko (2001) finds biases of the adjusted ISPI as ranging from –0.7% to 7.9% for the overall ISPI, from –3.2% to 3.6% for foods, from –4.4% to 9.0% for manufactured goods, and from –8.5% to 30.0% for services. Thus, noticeable biases can be accumulated even during one year with a 10% inflation.

3.2 Evolution of food price levels

Two sets of time series characterizing dynamics of prices for foods from price liberalization to present are available. The first is a sub-category of the CPI, the index of prices for foods. The second is the cost of a staples basket, which covers goods with the total weight in the Russian CPI exceeding half of the weight of foods. This makes it a good representative of the food basket involved in the CPI. We check whether a characterization of price behavior through price changes matches price behavior through absolute prices. Figure 3 compares these indicators with each other, demonstrating their trajectories by economic area over 1992-2005. In its upper panel, the trajectories of the ISPI-food are plotted. In the lower panel, the trajectories represent the basket-25, $P_{r0t}^{(25)}$, in 1992:2 through 2000:5, and the basket-33, $P_{r0t}^{(33)}$, since 2000:6.

The behavior of price levels in the upper and lower panels of Figure 3 is quite dissimilar. According to the ISPI, area food price levels diverged during 1992-1994, then almost stabilized, for the most part oscillating around permanent values. The data on the costs of staples suggest three stages in the evolution of the food price levels. The first stage is also divergence in 1992-1993 and, to some extent, in 1994. In contrast to the behavior of the ISPI, however, convergence of price levels started in 1994. This stage lasted until approximately the end of 1999. Since then, food price levels have been more or less stable.

Spatial features of the price levels are also different. Judging from the ISPI, the range of the levels in recent years is about from 65% to 150% of the national average. There is a cluster of five economic areas with close price levels (approximately 10% less than the national level), while the others differ significantly from one another in the price levels. According to the cost of staples, all areas but the Far East are concentrated in recent years in the band of ±10% around the national level. Thus, the range of price levels is about 90% to 110% of the national average. Considering the Far East, the upper bound of the range becomes, on average, about 130%.

Based on the fact of good consistency of the relative cost of staples basket with the Surinov DSPI, it can be believed that the former rather well approximates the dynamic behavior of the latter. Hence, Figure 3 provides additional evidence of significant biases in ISPI that severely distort spatial comparisons, even in a qualitative sense.

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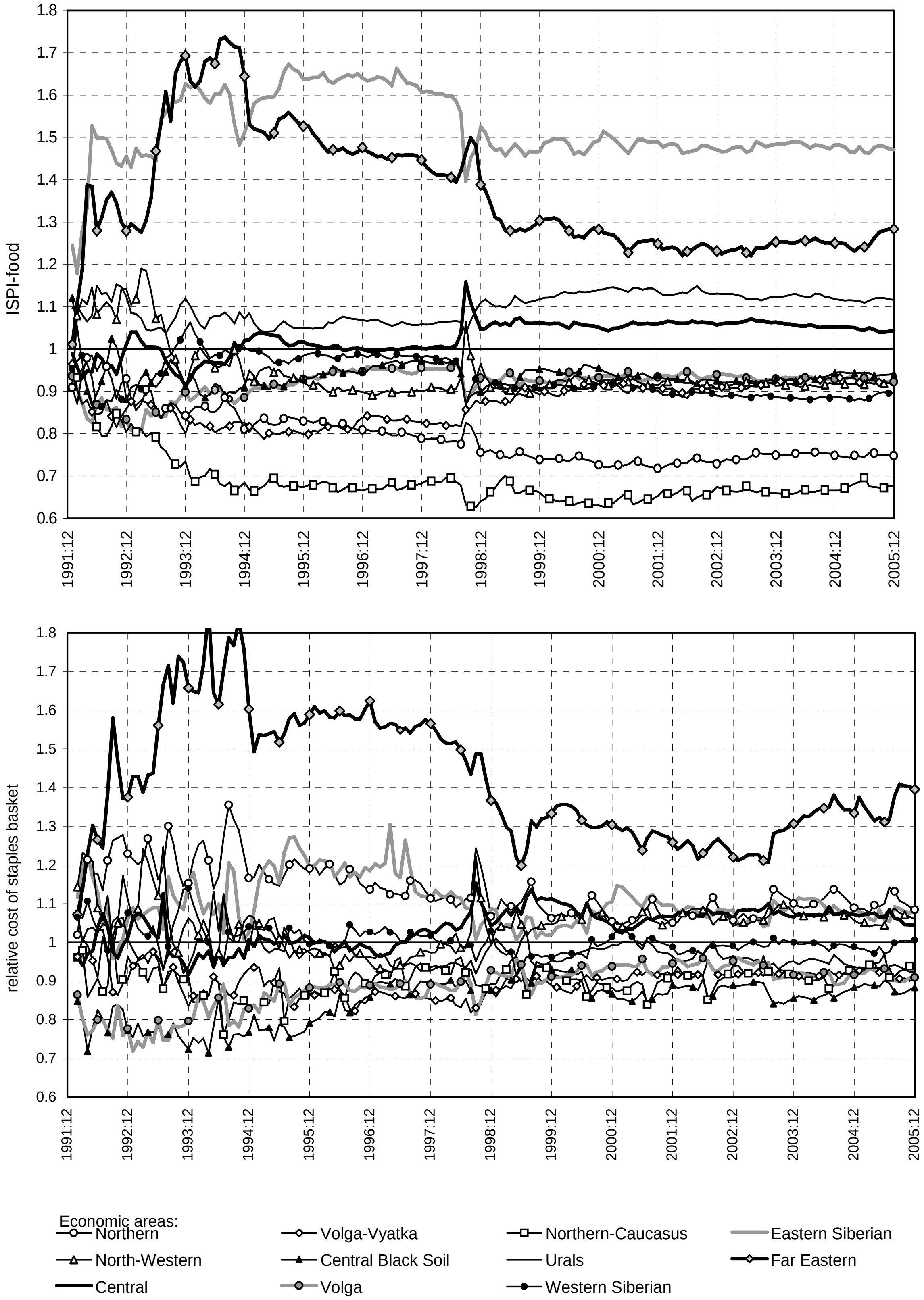
Figure 3. Dynamics of food price levels.



3.3 Indirect SPI vs. the purchasing power basket SPI

In this section, ISPIs are compared with the PP basket DSPI, $P_{r0t}^{(\mathrm{PP})}$. These are the ISPI defined by (7), $P_{r0t}^{(\pi)}$, and the ISPIs adjusted for biases accumulated by 1998:1 or 2002:1. To adjust an ISPI for the 1998:1 bias, the Surinov DSPI for 1998:1 is used as the base DSPI, i.e. $P_{r0t}^{\prime\prime(\pi)} = P_{r0,1998:1}^{(\mathrm{S})} \cdot I_{r0}(1998:2,t)$. The use of the cost of the PP basket in 2002:1 as the base DSPI, $P_{r0t}^{\prime\prime\prime(\pi)} = P_{r0,2002:1}^{(\mathrm{PP})} \cdot I_{r0}(2002:2,t)$, adjusts an ISPI for the 2002:1 bias. Table 7 reports results for two points in time: 2002:1, the earliest period, for which the costs of the PP basket are available, and for 2005:12, a recent period. Figures 4 and 5 compare the dynamics of $P_{r0t}^{(\pi)}$ and $P_{r0t}^{\prime\prime(\pi)}$, respectively, with that of $P_{r0t}^{(\mathrm{PP})}$ over the whole time span of 2002:1 through 2005:12.

Table 7. Comparison of ISPI with purchasing power basket DSPI

Economic area	**2002:1**			**2005:12**			
	ISPI	**ISPI adjusted for the 1998:1 bias**	**PP basket DSPI**	**ISPI**	**ISPI adjusted for the 1998:1 bias**	**ISPI adjusted for the 2002:1 bias**	**PP basket DSPI**
Northern	0.86	0.97	1.04	0.87	0.98	1.05	1.08
North-Western	0.83	1.01	1.02	0.83	1.01	1.02	1.02
Central	1.00	1.00	1.10	1.00	1.00	1.10	1.12
Volga-Vyatka	0.87	1.02	0.89	0.89	1.05	0.91	0.90
Central Black Soil	0.97	0.89	0.86	0.96	0.88	0.85	0.88
Volga	0.87	0.94	0.92	0.87	0.95	0.93	0.91
Northern-Caucasus	0.92	0.93	0.92	0.92	0.94	0.92	0.92
Urals	1.04	1.08	0.94	1.03	1.06	0.92	0.90
Western Siberian	1.12	0.93	1.00	1.10	0.91	0.99	0.99
Eastern Siberian	1.04	0.97	1.04	1.05	0.97	1.04	1.01
Far Eastern	1.40	1.22	1.29	1.43	1.24	1.31	1.35

Table 7 and Figures 4 and 5 suggest great quantitative discrepancies between ISPIs and the PP basket DSPI. Worse, the ISPIs yield patterns that are quite different qualitatively from that yielded by the PP basket DSPI. Even adjustment of the ISPI for the 1998:1 biases does not improve the situation. This raises the question as to whether the PP basket DSPI is consistent with the Surinov DSPI. If that is not the case, then it is the inconsistency between the two kinds of DSPIs that could be a reason behind the biases of adjusted ISPI relative to the PP basket DSPI.

Figure 4. Dynamics of the ISPI and the PP basket DSPI

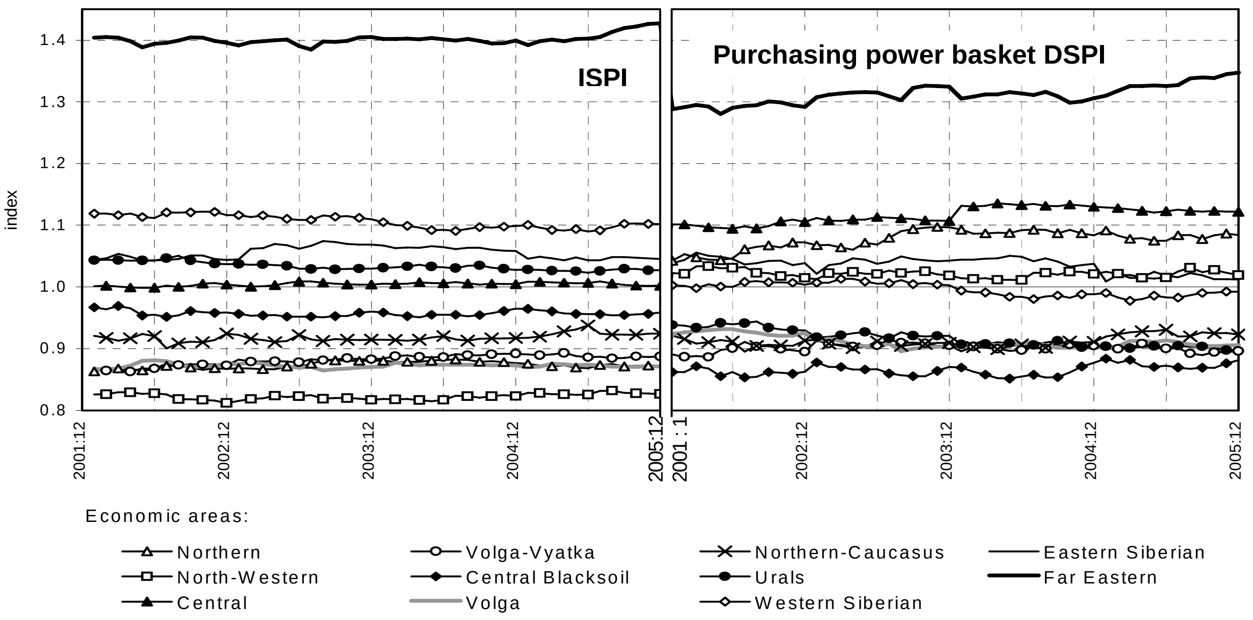
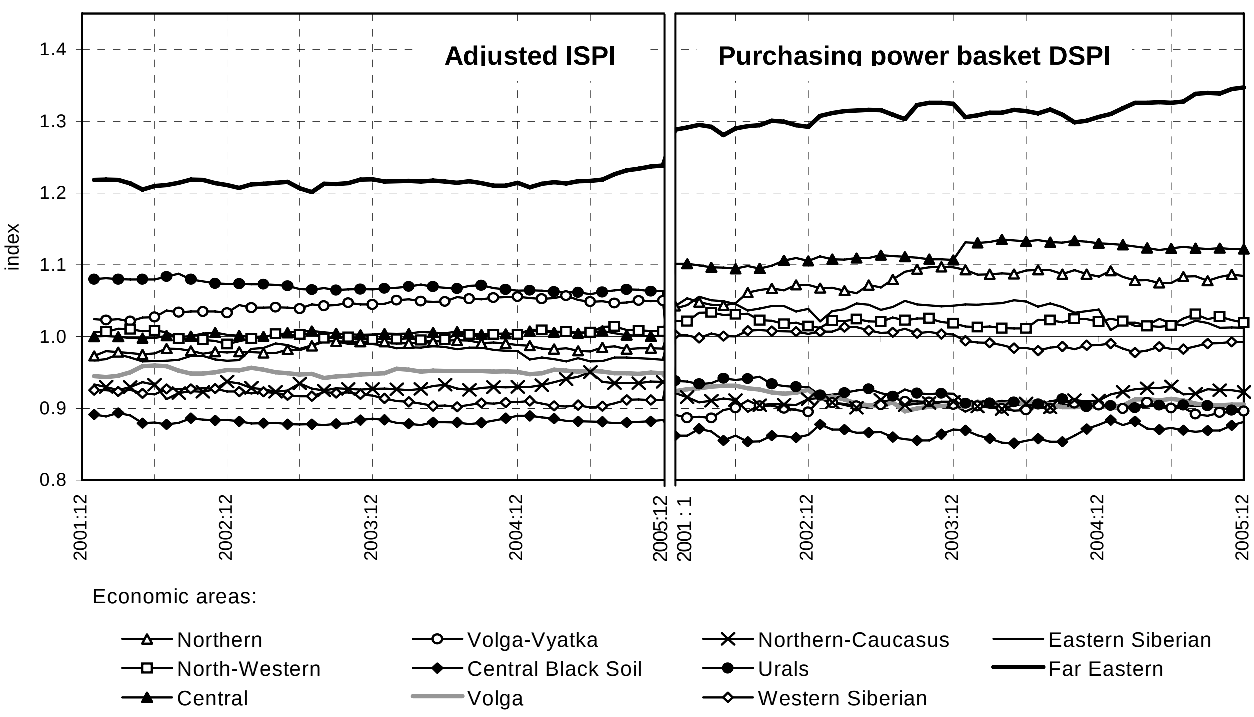


Figure 5. Dynamics of the ISPI adjusted for 1998:1 bias and the PP basket DSPI



There is no way to directly compare the Surinov and PP basket DSPIs, nor is a direct comparison of the relative cost of staples basket with the PP basket DSPI possible as the subcategories of the latter (in particular, the food component) are not published. Nonetheless, some indirect evidence can be exploited. The lower panel of Figure 3 suggests no fundamental changes in the ranking of economic areas by the cost of staples over 1998:1 through 2002:1. Based on this, the ranking by the overall price level in 2002:1 is essentially the same as in 1998:1. Thus, if the Surinov and PP basket DSPIs are consistent, the rankings by the former for 1998:1 and by the latter for 2002:1 should be fairly similar.

Indeed, excluding the Central economic area, these rankings prove close. The ranks differ for three areas only; for two areas the difference equals 1, and for the other, the North-Western economic area, the difference is 2. This can be caused by actual changes in the region ranks over the four years and/or by differences in the construction of the two DSPIs, which inevitably have an effect for locations with close values in either index. The special case is that of the Central economic area: its rank differs by 8 in the two rankings. Again, the behavior of prices for staples provides an explanation. Before the August 1998 financial crisis, the cost of the basket in the Central economic area was about the national average. After the crisis, the basket cost rose by about 10%, increasing the overall price level and the rank of the area. The relative cost of the staples basket was 1.025 in 1998:1, and 1.074 in 2002:1.

These considerations suggest that the Surinov and PP basket DSPIs are consistent, or at least not contradictory. Hence, the biases of the adjusted ISPI are caused by its own properties. Table 8 summarizes the biases in percentage terms.

Table 8. Percentage of ISPI biases relative to the purchasing power basket DSPI

Economic area	2002:1		2005:12		
	ISPI	ISPI adjusted for the 1998:1 bias	ISPI	ISPI adjusted for the 1998:1 bias	ISPI adjusted for the 2002:1 bias
Northern	-17.2	-6.7	-19.6	-9.3	-2.8
North-Western	-19.3	-1.6	-18.9	-1.2	0.4
Central	-9.1	-9.2	-10.7	-10.8	-1.7
Volga-Vyatka	-2.8	15.0	-1.0	17.2	1.9
Central Black Soil	12.1	3.4	8.8	0.3	-3.0
Volga	-6.1	2.3	-3.8	4.8	2.5
Northern-Caucasus	-0.1	1.3	0.1	1.5	0.2
Urals	11.2	15.1	14.3	18.3	2.8
Western Siberian	11.6	-7.7	11.1	-8.1	-0.5
Eastern Siberian	0.5	-6.9	3.3	-4.4	2.7
Far Eastern	9.0	-5.4	6.0	-8.1	-2.8

Figure 6 shows the pattern of biases in detail. It reports kernel estimates of empirical distributions of biases across regions. Table 9 tabulates summary statistics of these distributions.

Figure 6. Cross-region empirical distributions of ISPI biases relative to the PP basket DSPI

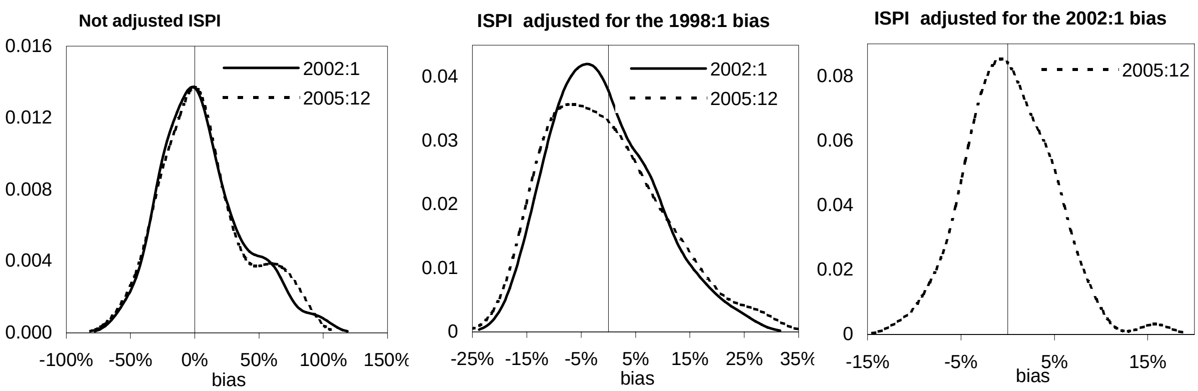


Table 9. Summary statistics for cross-region bias distributions (%)

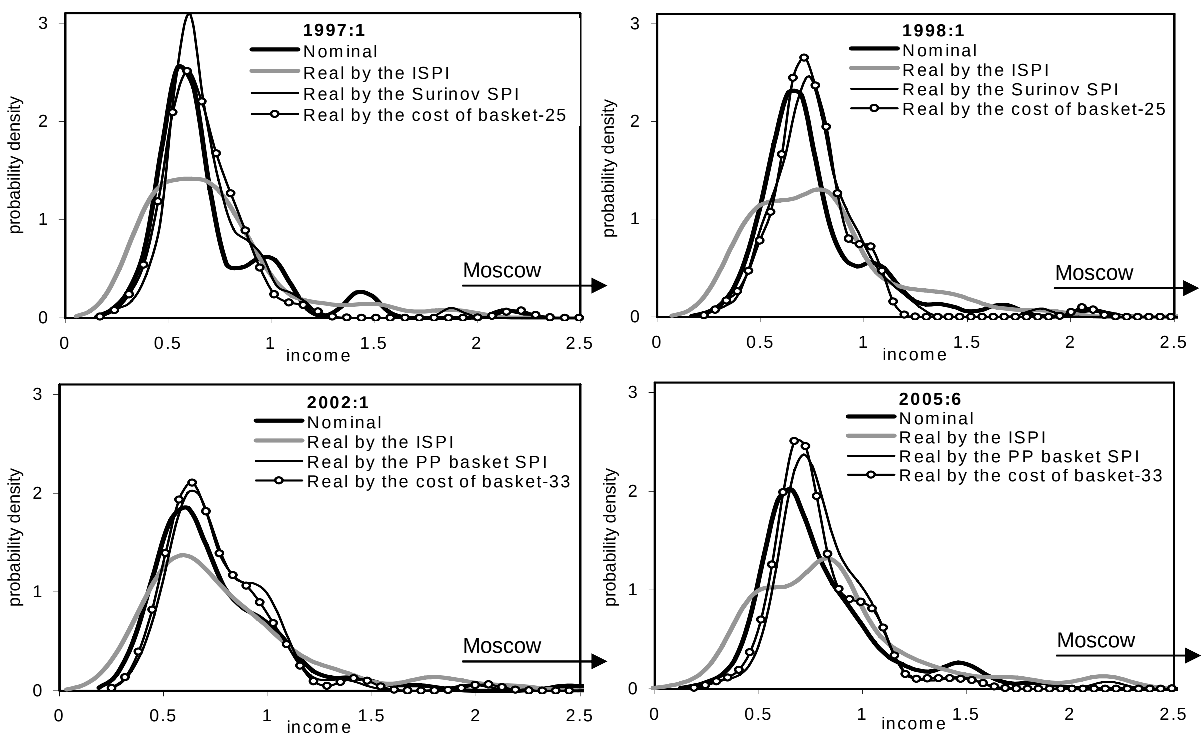
Adjustment of ISPI	Mean	Median	Standard deviation	Minimum	Maximum
2002:1					
No adjustment	6.5	1.1	32.4	-57.0	98.1
For the 1998:1 bias	-0.8	-1.4	9.1	-16.9	24.7
2005:12					
No adjustment	6.8	1.3	33.0	-57.9	85.5
For the 1998:1 bias	-0.6	-2.3	10.4	-17.1	27.7
For the 2002:1 bias	0.2	0.1	4.6	-11.2	15.8

A feature evident in Tables 8 and 9 (as well as in Table 3 in Section 3.1) is that biases tend to increase over time. The variance of biases rises and the bias range, as a rule, widens when we compare an ISPI, not adjusted or adjusted for biases, with itself after some time. Of course, the bias in a particular location may either increase or decrease over time. Consider, for example, the 1998:1-adjusted ISPI in Table 8; its bias ranges from –9% to 15% in 2002, and from –11% to 18% in 2005. The average (across economic areas) absolute bias rises from 6.8% in 2002:1 to 7.6% in 2005:12.

4 Estimating real incomes

Having found sufficient biases in the indirect SPI, we now attempt to gain insight into their effect for spatial comparisons of economic indicators expressed in real terms by estimating per capita income distribution in Russia. The raw income data are drawn from Goskomstat/Rosstat's monthly bulletins *Socio-Economic Situation of Russia*. In this section, data across regions are exploited rather than across economic areas. We estimate distributions for four points in time: 1997:1 and 1998:1, for which the Surinov DSPI is available; and 2002:1 and 2005:6, when the PP basket DSPI is available.

Figure 7. Empirical distributions of regional incomes per capita in Russia



In the following discussion, incomes are normalized to the national average. Kernel estimates of the densities of nominal and real incomes are shown in Figure 7.

Three versions of real income are compared. The first is an estimate through the ISPI: $M_{rt}^{(\pi)} = M_{rt} / P_{r0t}^{(\pi)}$, where M_{rt} is the normalized nominal per capita income in region r at time t. The second is an estimate through a DSPI; for 1997 and 1998, the Surinov DSPI is used, $M_{rt}^{(S)} = M_{rt} / P_{r0t}^{(S)}$, and for 2002 and 2005, the PP basket DSPI is used,

$M_{rt}^{(\mathrm{PP})} = M_{rt} / P_{r0t}^{(\mathrm{PP})}$. The third is real income estimated through the cost of a staples basket. We use the basket-25 for 1997 and 1998 ($M_{rt}^{(25)} = M_{rt} / P_{r0t}^{(25)}$), and the basket-33 for 2002 and 2005 ($M_{rt}^{(33)} = M_{rt} / P_{r0t}^{(33)}$). Table 10 reports summary statistics for income distributions.

Table 10. Summary statistics for income distributions

Version of real income estimate	**Mean**	**Median**	**Standard deviation**	**Minimum**	**Maximum**		
					Without outliers	**Tyumen Oblast**	**Moscow**
			1997:1				
Nominal (*M*)	0.743	0.615	0.472	0.282	1.486	2.149	3.880
M/ISPI	0.741	0.655	0.443	0.252	1.936[a]	1.794	3.277
M/Surinov SPI	0.720	0.631	0.390	0.282	1.150	1.880	3.531
M/cost of basket-25	0.713	0.632	0.403	0.307	1.166	2.189	3.488
			1998:1				
Nominal (*M*)	0.806	0.687	0.413	0.317	1.699	2.097	3.278
M/ISPI	0.810	0.741	0.400	0.273	1.907[a]	1.722	2.724
M/Surinov SPI	0.797	0.746	0.332	0.324	1.230	1.853	2.983
M/cost of basket-25	0.780	0.724	0.320	0.334	1.166	2.067	2.756
			2002:1				
Nominal (*M*)	0.802	0.666	0.511	0.381	1.723	2.457	4.198
M/ISPI	0.822	0.701	0.453	0.267	1.967	2.182	2.809
M/PP basket SPI	0.794	0.710	0.330	0.413	1.393	1.976	2.697
M/cost of basket-33	0.785	0.698	0.384	0.410	1.444	2.037	3.259
			2005:6				
Nominal (*M*)	0.830	0.695	0.404	0.292	1.614	1.843	3.229
M/ISPI	0.864	0.836	0.416	0.190	2.238[b]	1.711	2.146
M/PP basket SPI	0.824	0.766	0.253	0.336	1.418	1.515	2.183
M/cost of basket-33	0.808	0.729	0.298	0.331	1.401	1.526	2.671

[a] Vologda Oblast
[b] Republic of Yakutia

Note that the national average is not an arithmetic mean over regions. It is the total income in the country divided by its population (i.e. the average over regions weighted by the regional share of Russia's population). That is why the main modes and means of the distributions are not close to 1 representing the national mean. The two obvious outliers are incomes in Moscow and the Tyumen Oblast. Small local modes corresponding Moscow are left beyond the plots in Figure 7, since they lie far from the general body of the observations.

Figure 7 suggests that when real incomes are estimated through the ISPI, the shape of the income distribution is severely distorted as compared to the use of direct SPIs. Even nominal income seems a better proxy of real income than $M_{rt}^{(\pi)}$.

Quantitatively (see Table 10), $M_{rt}^{(\pi)}$ shows wider variation across regions than other estimates, both in terms of range and standard deviation. For all four points in time, minimal real income (left bound of the income range) turns out to be smaller, while maximal real income excluding outliers (right bound of the range) turns out to be greater than those estimated through other SPIs. A striking feature is that sometimes the right bound of the range of $M_{rt}^{(\pi)}$ overlaps the outliers (even Moscow in 2005:6).

As they are monthly (instantaneous), the above results may suffer from random shocks and obscure a true pattern. To check for robustness, incomes are also estimated over 12-month time spans using geometric averaging. Because we lack time series for the Surinov DSPI, the same value is applied for all months of a relevant time span. The results appear in Appendix A, Figure A1 and Table A1. A comparison of these results with those in Figure 7 and Table 10 reveals no significant differences.

It follows from comparison of $M_{rt}^{(\mathrm{S})}$ with $M_{rt}^{(25)}$ and $M_{rt}^{(\mathrm{PP})}$ with $M_{rt}^{(33)}$ that the use of the cost of staples basket for estimating real incomes provides a rather good approximation of the income distributions based on the DSPIs. One possible explanation is that the share of foods in the total consumers' expenditures remained more or less stable before the 1998 financial crisis (Goskomstat, 2001, p. 167). Spending patterns on food change in the 2000s as Russia begins to experience an economic upturn. As a result, the quality of approximation worsens in 2002 and 2005 as compared with 1997 and 1998. In any case, the cost of staples basket remains a better proxy of regional price levels than the ISPI from the viewpoint of the shape of real income distribution.[14]

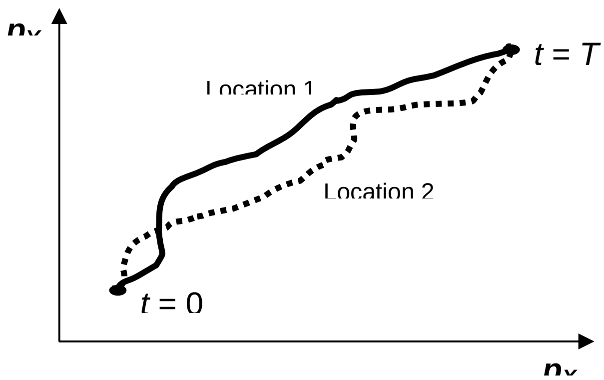
While the pattern based on real incomes estimated by DSPIs demonstrates that differences in regional costs of living smooth to some extent income inequality across Russian regions, estimates through the ISPI suggest the opposite – that inequality in real incomes exceeds inequality in nominal incomes. Thus, there is a danger that ISPI distortions could lead researchers and policymakers to inappropriate conclusions as to income convergence and income mobility of Russian regions.

5 Numerical experiments

5.1 Description

In this section, simulated data are used to understand the extent of ISPI biases under favorable (and controlled) conditions. In our numerical experiments, we consider a time span $t = 0,\dots,T$ and two locations. At starting point $t = 0$, prices are equal both across locations and goods. They randomly change but eventually return to equal values at final point $t = T$ as displayed in Figure 8. Comparing DSPI and ISPI at T, a bias of ISPI is estimated. Representative consumers are assumed to be identical across locations; that is, they have the same preferences and nominal incomes (valid for all $t \in [0,\dots,T]$). Under such assumptions, the DSPI equals 1 at time points 0 and T under any definition of DSPI. The ISPI is computed by Formula (2) with the base DSPI equaling 1 and the CPI being the Divisia index. Generating a great set of random price paths, a distribution of the ISPI bias is obtained. Under these conditions, the sole source of bias is the path-dependence of CPI. Two goods are designated X and Y (say, X for foods and Y for non-foods), and two locations are designated 1 and 2. The indexes would then be $k = X, Y$ and $r = 1, 2$, respectively.

Figure 8. Price paths.



Modeling consumption. Three demand systems model consumer behavior. Under $M / p_X \geq q_{X0}$, they look like:

$$\begin{cases} q_X = X_0 + \alpha(M - q_{X0}\, p_X)/p_X \\ q_Y = (1-\alpha)(M - q_{X0}\, p_X)/p_Y \end{cases}; \qquad (10)$$

[14] This makes Yemtsov's (2005) approach, mentioned in section 2.3, to appear quite reasonable.

$$\begin{cases} q_X = q_{X0}(1+\ln\dfrac{M}{q_{X0}p_X}) \\ q_Y = (M - q_{X0}(1+\ln\dfrac{M}{q_{X0}p_X})p_X)/p_Y \end{cases}; \tag{11}$$

$$\begin{cases} q_X = \beta q_{X0}(1-e^{-\delta M/q_{X0}p_X}) \\ q_Y = (M - \beta q_{X0}(1-e^{-\delta M/q_{X0}p_X})p_X)/p_Y \end{cases}, \quad \delta = \ln\frac{\beta}{\beta-1}. \tag{12}$$

Under $M/p_X < q_{X0}$, all the three take the form $q_X = M/p_X$, $q_Y = 0$, implying a corner consumer equilibrium. In the above formulas, q_X and q_Y are quantities, p_X and p_Y are prices, M is income, q_{X0} is a component of the subsistence minimum taken to be $q_X = q_{X0}$ and $q_Y = 0$, and α and β are parameters; $\alpha < 1$; $\beta > 1$. To economize notations, the location and time subscripts for quantities and prices are suppressed.

All the three demand systems are derived from nonhomothetic preferences. System (10) is yielded by Stone-Geary preferences $U(q_X, q_Y) = (q_X - q_{X0})^{\alpha} q_Y^{1-\alpha}$. System (11) implies that the income elasticity of demand for X asymptotically tends to zero with increasing quantity of X: $\varepsilon_{IX} = q_{X0}/q_X$, with $U(q_X, q_Y) = q_Y \exp(\int \frac{dq_X}{q_{X0}e^{q_X/q_{X0}-1} - q_X})$ as the relevant preferences. In system (12), consumption of X is assumed to have an upper limit βq_{X0}, approached as $M/p_X \to \infty$. Preferences

$U(q_X, q_Y) = q_Y \exp(-\int \frac{dq_X}{q_{X0}\ln(1-q_X/\beta q_{X0})/\delta + q_X})$ yield such a demand system.

Simulating inflation. Changes in prices over $[t–1, t]$, π_{rkt}, are generated randomly, differing across locations and goods (and certainly across time). They are normalized so that all prices are uniform at the final time point: $p_{krT} = p_{kr0}(1+\bar{\pi})^T$ for all r and k, where $\bar{\pi}$ is a given average inflation rate. That is, with ν_{krt} as the "raw" changes in prices, $\pi_{krt} = (1 + \nu_{krt})/a_{kr} – 1$. The normalizing factor is $a_{kr} = \prod_{t=1}^{T}\left(\frac{1+\nu_{krt}}{1+\bar{\pi}}\right)^{1/T}$. In turn, ν_{krt} is yielded by an autoregressive process AR(1): $\nu_{krt} = \rho\nu_{kr,t–1} + \varepsilon_{krt}$, $\nu_{kr0} = \varepsilon_{kr0}$, where ρ is an autoregressive coefficient ($0 \le \rho \le 1$). To make price cutting less likely than a rise in prices, shocks ε_{krt} are drawn from an asymmetric distribution: $\varepsilon_{krt} < 0$ are drawn from $N(0, \sigma_–^2)$ with probability $\sigma_–/(\sigma_– + \sigma_+)$, and $\varepsilon_{krt} \ge 0$ are drawn from $N(0, \sigma_+^2)$ with probability $\sigma_+/(\sigma_– + \sigma_+)$; $\sigma_–$ and σ_+

are standard deviations of negative and positive shocks; $\sigma_- < \sigma_+$. Prices are assumed to change continuously. Within intervals $[t–1, t]$, the changes are linear: $p_{kr}(t-1+\tau) = p_{kr,t-1}\cdot(1+\pi_{krt}\tau)$, $\tau \in [0, 1]$.

Incomes. The total expenditure on all commodities equals nominal income: $q_{Xrt}p_{Xrt} + q_{Yrt}p_{Yrt} = M_{rt}$. In contrast to prices, incomes, when not held constant, change discretely and remain constant within intervals $(t–1, t]$. That is, $M_r(t-1+\tau) = M_{rt}$, $\tau \in (0, 1]$. As mentioned, $M_{1t} = M_{2t}$ for all t.

Computing ISPI. A CPI over $[0,\ldots,T]$ for location r is computed as the Divisia index

$$I_r(0,T) = \exp\left(\int_0^T \frac{q_{Xr}(t)\cdot dp(t)_{Xr}/dt + q_{Yr}(t)\cdot dp_{Yr}(t)/dt}{M_r(t)}dt\right). \tag{13}$$

Given that prices are piecewise-linear functions of time and expenditures are piecewise constant, (13) takes the form

$$\begin{aligned} I_r(0,T) = \exp(\sum_{t=1}^{T}(\frac{p_{Xr,t-1}\pi_{Xrt}}{M_{rt}}\int_0^1 q_{Xr}(p_{Xr}(t-1+\tau), p_{Yr}(t-1+\tau))d\tau + \\ \frac{p_{Yr,t-1}\pi_{Yrt}}{M_{rt}}\int_0^1 q_{Yr}(p_{Xr}(t-1+\tau), p_{Yr}(t-1+\tau))d\tau)) \end{aligned}, \tag{14}$$

where $p_{kr,t-1} = p_{kr0}\prod_{z=1}^{t-1}(1+\pi_{krz})$. To compute (14), numerical integration is implemented.

Since the DSPI at $t = 0$ equals unity, the ISPI at $t = T$ is $P_{12T}^{(\pi)} = I_{1T}(0,T)/I_{2T}(0,T)$, being referred to as the Divisia ISPI. Its deviation from unity characterizes the bias relative to the DSPI.

For comparison, the "Young ISPI" is also computed, using a chained Young index similar to (3) and (4) with yearly updated weights. For two goods, (4) looks like

$$I_r(t-1,t) = w_{Xr\theta}\frac{p_{Xrt}}{p_{Xr,t-1}} + w_{Yr\theta}\frac{p_{Yrt}}{p_{Yr,t-1}}, \tag{15}$$

where $\theta = [(t-1)/12]$, $[x]$ is integer part of x;

$w_{Xr\theta} = \sum_{t=1}^{12} q_{Xr,(\theta-1)\cdot 12+t}\,p_{Xr,(\theta-1)\cdot 12+t} / \sum_{t=1}^{12} M_{r,(\theta-1)\cdot 12+t}$ for $\theta \geq 1$, and $w_{Xr0} = q_{Xr0}p_{Xr0}/M_{r0}$ for $\theta = 0$; $w_{Yr\theta} = 1 - w_{Xr\theta}$.

5.2 Results

The results reported below are obtained for T = 120 (10 "years" × 12 "months"). The average monthly inflation rate, $\bar{\pi}$, equals 1.35%, yielding a fivefold rise in prices over the whole time span, $p_{krT} = (1.0135)^{120} \approx 5p_{kr0}$; $\sigma_{-} = \bar{\pi}/2$ and $\sigma_{+} = \bar{\pi}$; ρ = 0.7. Figure 9 depicts a kernel estimate of the distribution of simulated inflation rates π_{krt}.

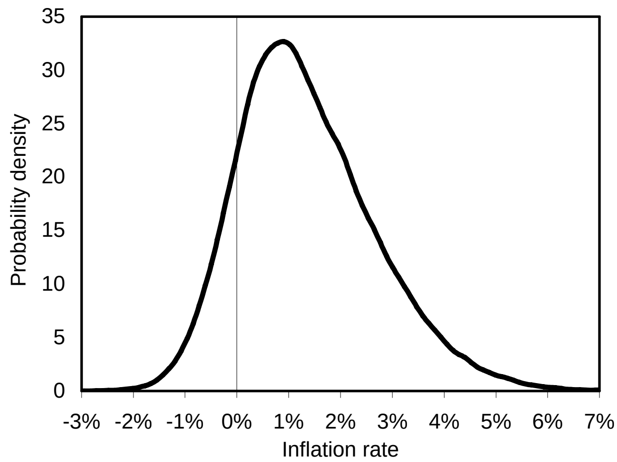


Figure 9. Distribution of simulated inflation rates.
(mean = 0.0136; median = 0.0119; min = –0.0283; max =0.0826; standard deviation = 0.0133).

The number of replications is 10,000 in each experiment. Parameters of the demand systems are: α = 0.3, β = 1.5, and q_{X0} = 0.9. Starting prices are p_{rk0} = 1. Nominal incomes are set in two ways. The first is to hold incomes constant during [0,…,T], whereby M_{rt} is set equal to 5.5. The second, more realistic, approach assumes that real incomes will rise by the end of a ten-year span of time. Nominal incomes are assumed to steadily rise with a constant monthly rate $(1+\bar{\pi})m^{1/T}-1$, so that the real incomes at the final point $t = T$ become m times higher than at the starting point; m is taken to equal 2 with starting incomes M_{r0} = 1. A bias of ISPI is computed as (ISPI – DSPI)/DSPI, i.e., ISPI – 1 (since DSPI = 1). Figure 10 demonstrates kernel estimates of the ISPI bias distributions obtained, and Table 11 reports summary statistics for the distributions.

Figure 10. Distributions of simulated ISPI biases

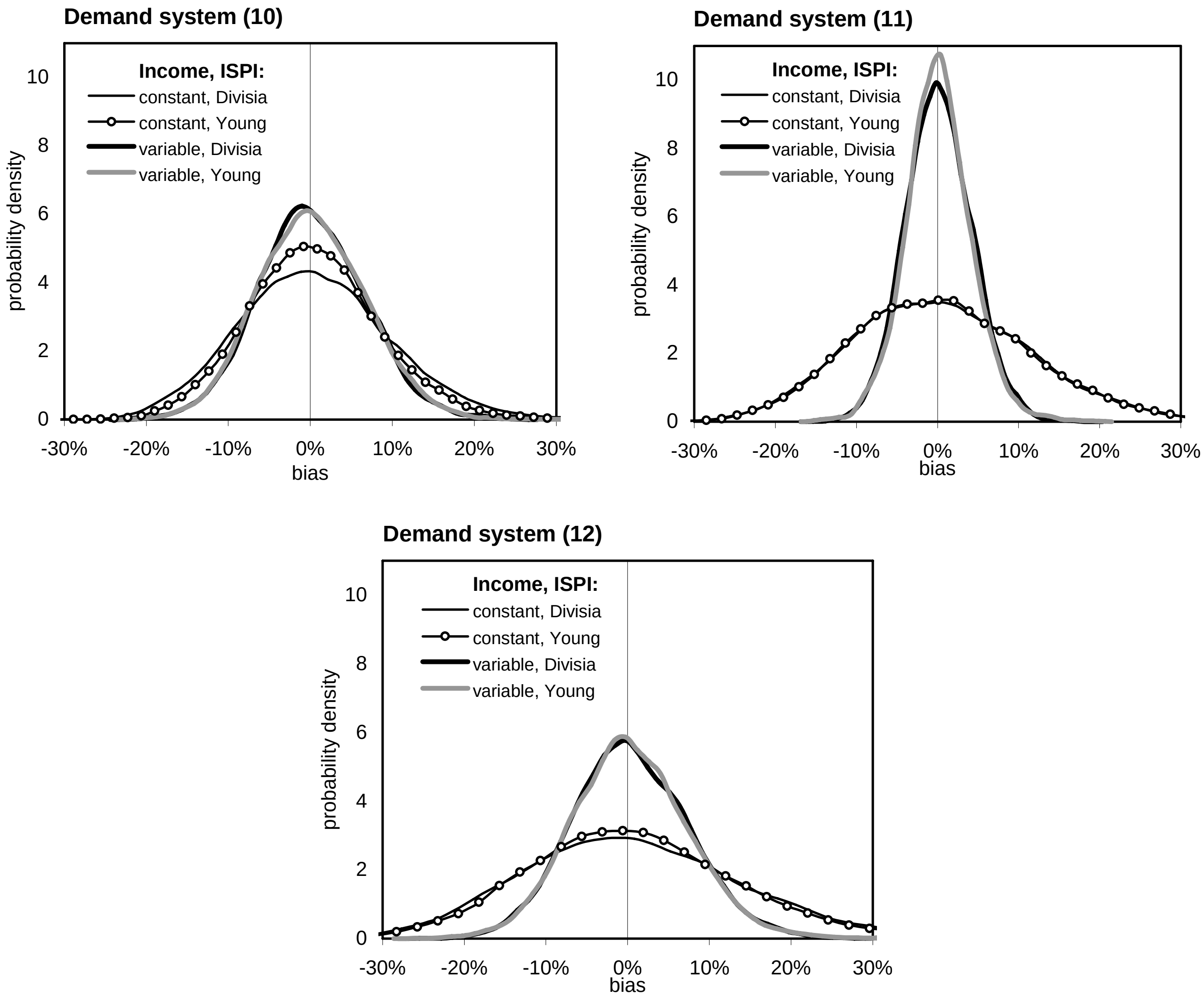


Table 11. Summary statistics for ISPI bias distributions (%)

Demand system	Nominal income	ISPI	Mean	Minimum	Maximum	Standard deviation
(10)	constant	Divisia	0.5	–31.7	44.4	9.3
		Young	0.4	–31.6	44.8	8.1
	variable	Divisia	0.2	–22.6	25.8	6.5
		Young	0.3	–21.5	31.0	6.6
(11)	constant	Divisia	0.6	–35.0	44.6	11.2
		Young	0.6	–36.8	50.3	11.1
	variable	Divisia	0.2	–15.7	19.2	4.1
		Young	0.2	–15.7	20.4	4.1
(12)	constant	Divisia	1.0	–41.3	58.4	14.1
		Young	0.9	–42.3	73.8	13.2
	variable	Divisia	0.4	–26.1	35.0	7.1
		Young	0.4	–26.7	39.0	7.2

These results indicate that only the path-dependence of CPI can sufficiently bias ISPIs. The most impressive are ranges of biases in Table 11, suggesting that the use of ISPI might understate SPI by over 40% or overstate it by over 70%. Dispersion of biases (measured by standard deviations) is large as well, varying across experiments from 4% to 14%. We see that holding nominal income constant, the dispersion increases as compared to a respective experiment with variable income. The reason is a difference in changes of real income. In the case of constant nominal income, the real income changes (increases) fivefold by the end of the time span, while changing (decreasing) only twofold in the case of variable nominal income.

Figure 10 suggests that the distributions of biases are nearly symmetric around zero. Hence, estimates of SPI by ISPI can be either understated or overstated with approximately equal probability. With the exception of demand system (11) under variable incomes, the shapes of the distributions are roughly similar across demand systems. This provides hope that the pattern is qualitatively similar to what is actually occurring in the real world, whatever a real demand system may be.

Although we know the Young index is biased compared to the Divisia index, the distributions of the Divisia and Young ISPI biases are surprisingly close to each other. The explanation is that these biases are approximately equal for locations 1 and 2. Therefore, they almost cancel out in the ISPI which is the ratio of location CPIs. As a digression, it seems interesting to consider the biases of the Young index relative to the Divisia index.

Appendix B provides data on the distributions of such biases in the above experiments. Despite the widely held belief that the Young index overstates inflation, the probability of downward biases surprisingly proves much higher than that of upward biases. The point is that this belief is true only for homothetic preferences. For nonhomothetic preferences, the Young index can deviate from the Divisia index in either direction depending on properties of specific preferences.[15]

The experiments not reported here may be summarized as follows. First, the higher and the more volatile inflation, the greater biases of ISPI (i.e. the standard deviation of their distribution). This is valid for increases in both the average inflation rate, $\bar{\pi}$, and cumulative inflation with widening the time horizon T at a fixed $\bar{\pi}$. Volatility of inflation rises with increasing $\sigma_{-} + \sigma_{+}$ and/or ρ. Second, the more deviation of preferences that generate a given demand system from homotheticity, the greater the biases. Such a deviation is determined by parameters q_{X0} and β. Finally, the lower income level and the greater change in real income over time, the greater the biases. Given M_{r0}, the standard deviation of biases is minimal when the nominal income increases with monthly rate $\bar{\pi}$ so that the real income remains constant on average. In this case, the standard deviation further diminishes with increasing M_{r0}.

6 Reasons for the discrepancy between indirect and direct SPIs

There are three types of reasons for biases of ISPIs relative to DSPIs found in Section 3. The first consists of theoretical reasons that cause fundamentally unavoidable biases. The second group consists of practical reasons. As approximations of a true estimator of price change (e.g. the Divisia index), the CPIs as computed give rise to additional biases. The third type involves Russia-specific features of the empirical data used in this study.

Theoretical reasons. As mentioned in the introduction, there is a conceptual inconsistency between price levels involved in comparisons across space and across time. The Konüs

[15] Taking $q_{X0} = 0$ in (10) to get the Cobb-Douglas preferences, we find *no* downward bias of the Young index in the bias distribution.

(1924) theory of the cost-of-living index might provide a conceptual framework to consider this issue, measuring the absolute price level in a location, P_r, as the minimum cost of achieving a given utility level. However, for simplicity's sake, P_r is assumed to be measured as the cost of a fixed commodity basket. The reasoning below can then be easily reinterpreted in the spirit of the Konüs theory. The basket represents the consumption pattern of local consumers and is defined as a set (a vector) of quantities of commodities $\{k\}$: $\mathbf{q}_r = (q_{kr})$; with $\mathbf{p}_r$ as the price vector, $P_r = \mathbf{q}_r\mathbf{p}_r$.

Let prices in locations r and s be equal, $\mathbf{p}_r = \mathbf{p}_s$. Then it is reasonable to deem the SPIs for these locations equal to one: $P_{rs} \equiv P_r/P_s = P_{sr} \equiv P_s/P_r = 1$. In doing so, the location baskets are implicitly assumed to be identical, $\mathbf{q}_r = \mathbf{q}_s = \mathbf{q}$. Put differently, it is assumed that there is a single representative consumer (with consumption pattern $\mathbf{q}$), who confronts the price vector at two locations. CPI baskets generally are not identical across the locations, however, so even if consumers representing location have identical preferences, their consumption patterns will differ (at a weight reference period) due to unequal incomes. If incomes are equal, but preferences are different, the differences in consumption patters are likely even greater. Thus, equality of prices across locations does not imply equality of price levels as they are defined in the local CPIs.

The CPI is intended for comparing an individual local consumer's behavior to his behavior at another point in time. However, once the CPI is involved in estimation of indirect SPI, the comparison is made between different consumers. In other words, we assume considering a single representative consumer, when in fact we are dealing with two different consumers. In essence, the inconsistency of the ISPI is in that the cost of one commodity basket is taken to measure the base price level, while change in the cost is measured for a different basket. Moreover, the former basket is uniform across locations, and the latter is location-specific. Thus, it does not matter whether a particular DSPI basket is dealt with (as in the case where the Surinov DSPI or a basket DSPI is taken for the base DSPI) or whether the basket is not explicitly defined (as in the case of assuming equality of prices to imply the unity base DSPI). As empirical findings in Section 3 suggest, this inconsistency yields substantial ISPI biases in either case.

The situation is not improved by the application of more sophisticated DSPIs such as the Geary-Khamis index or the Éltetö-Köves-Szulc index for the base DSPI. Basically, the averaging used in these indexes is equivalent to construction of a representative consumer common for all locations involved in the comparison. However, the CPI in relevant ISPIs implies different, location-specific, representative consumers.

In principle, the DSPI can account for variability of consumption patterns (preferences) across locations due, say, to national tastes, climate, etc. In such a case, however, location-specific baskets should refer to the same utility level. An example is a subsistence minimum basket varying across locations. It is assumed that there is a uniform minimal level of welfare, and that such a level is provided in each location by its own commodity basket. Again, the CPI basket in any one location will differ from its subsistence minimum basket.

The second theoretical reason for bias in the ISPI is the path-dependence of CPI even if a "perfect" continuous time index number (Divisia index) is used. Hence, any chain CPI as its approximation will all the more be biased. As Samuelson and Swamy (1974) prove, the Divisia index is path-invariant if and only if preferences are homothetic. This case is highly unrealistic as it implies unity income elasticity of demand for all commodities.

As Section 5 demonstrates, the path-dependence of the CPI can be a source of sufficient biases in the ISPI, distorting the measure of the overall price change between the beginning and the end of a time span under consideration. In our numerical experiments, the mean of biases is close to zero, and the distributions prove to be symmetric, which suggests that downward and upward biases have equal probability. However, this is due to the assumption of identical consumer preferences across locations. Such an assumption is hardly true in reality. Given spatially differing preferences, asymmetry of the bias distribution and nonzero mean is to be expected. In this case, bias in one direction or another would prevail, further distorting the pattern of spatial comparisons (e.g. by systematically overstating or understating real incomes).

Practical reasons. CPIs are computed with an approximating method, typically a Young ("Laspeyres-type", "modified Laspeyres") index. Young indexes are known to produce a number of biases, and particularly a substitution bias. Although this fact has long been explored theoretically, empirical examination of such bias in real CPIs started just over a decade ago. In their thorough analysis of the US CPI, Boskin *et al*. (1996) found a persistent upward bias of 1.1 percentage points per year.

Bessonov (1998) estimated the upper level substitution bias of the Russian national CPI. His results suggest that the official CPI overestimates the rise in prices in Russia by 35% over the 1992-1996 period. Even though the weights in the Russian CPI were updated every year, the frequency proved insufficient to keep pace with the dynamic changes in prices during periods of high inflation. Using a different methodology, Gibson, Stillman, and Le (2004) found the upward bias in the Russian national CPI to equal at least 30% over the 1992-2001 period.

Appendix B characterizes biases of the Young index (substitution biases) in artificial examples taken for the numerical experiments. These indicate that the bias distributions are asymmetric. However, the distributions of the Divisia and Young ISPI biases turned out to be close to each other (see Figure 10), which suggests that the substitution biases are approximately equal across locations, almost canceling out. Such a result is due to the identity of preferences and nominal incomes across locations in the experiments. Of course, this does not take place in reality. In particular, the standard deviation of nominal income per capita across Russian regions ranges up to more than a half of the national average (see Table 10). We should thus expect the substitution bias to be region-specific and not to cancel out in ISPIs. As a result, ISPIs contain a combination of the path-dependence bias and substitution bias.

Some price index formulas generate bias even when consumers are not substituting commodities. For example, such formulas can indicate inflation while it does not actually occur. Since the Young index is irreversible in time, it causes the chain index to increase while prices oscillate in some range. Seasonal oscillations (as in Russia, where inflation sometimes alternates with summer deflation) add spurious inflation to actual inflation in the value of the index. The magnitude of this artifact varies across locations, depending on location-specific amplitudes of oscillation. As Auer (2002) demonstrates, the Laspeyres index can also indicate spurious inflation, although this index is time-reversible. Auer's inversion test characterizes a situation where pairs of commodities display inverse developments. Such a situation can occur, for example, in the course of seasonal deflation with asynchronous price-cutting.

A disadvantage of the chain CPI is that it keeps errors made in its links and accumulates them over time. Once made, a measurement error remains in the chain forever, and further measurements do not correct it. Thus, an inadvertent error could permanently bias the CPI.

Russia-specific reasons. The assumption that the base DSPI equals one rests on a further assumption of uniformity of prices across Russian regions in December 1991. As discussed in Section 2.2, such a uniformity assumption is problematic. While cross-region variability of prices at the base period was probably not large, there is no way to determine the actual variation. Thus, it is not inconceivable that a starting bias may be embedded in the ISPIs analyzed.

Measurement errors are also quite likely in the early years of transition. Regular statistical observation of prices was introduced in contemporary Russia only in 1990. Thus, at the beginning of transition, Russian statisticians had no experience in price observation, let alone in a market environment. Moreover, extremely high inflation made tracking prices a non-trivial task. Errors made in the early 1990s are incorporated, as discussed above, in all further values of the Russian CPIs. It is measurement errors in 1992 that likely account for the incredibly high prices for foods in Eastern Siberia according to the ISPI (see upper panel of Figure 3).

The total bias of the ISPI is a superposition of the above particular biases. The use of the Surinov DSPI for 1997:1 or 1998:1, or the PP basket DSPI for 2002:1 as the base DSPI eliminates biases caused by the Russia-specific reasons, the (possible) starting bias and the measurement errors in the early years of transition. Nonetheless, the bias-adjusted ISPIs turn out to be biased again as well (see Tables 5, 6, 8, and 9 and Figure 5). The total ISPI bias is time-dependent. Moreover, the path-dependence bias and the substitution bias can change with time in either direction under nonhomothetic preferences. Therefore, the bias of ISPI for some location can increase over one time interval and decrease over another. Tables 2 and 6 suggest that this is the case. This makes the dynamic pattern yielded by some ISPI quite obscure, so we cannot conclude whether price levels, real incomes, etc. across regions actually tend to converge or diverge.

7 Conclusions

To provide spatial comparability, local monetary indicators should be normalized by local price levels. A widespread practice is evaluation of such levels indirectly using local CPIs. It is essential that the method applied provides a good approximation of actual price levels, or at least yields a qualitative pattern consistent with such price levels. Taking data across Russian regions, this paper demonstrated that such is not the case. The ISPIs are substantially biased and distort the qualitative pattern. In general, the ISPIs tend to overstate cross-spatial differences. In particular, ISPIs were shown to provide the worst conversion of nominal incomes across Russian regions into real incomes. Even nominal incomes turned out to be a better proxy of real incomes that those normalized by the ISPIs.

This sad story implies that there are no means to adequately estimate real incomes, real wages, real consumption, cost of living, etc. across regions of Russia before 2002, with the possible exception for the cusp of 1997/1998. Indeed, the only item offering a potentially more or less adequate approximate of regional price levels before 2002 is the cost of a staples basket.

While the empirical data used are those for Russia, the general conclusion here applies for other countries as well. There are only two potential Russia-specific reasons for biases and these were eliminated in some of our results. Moreover, the Russian CPI is conceptually quite similar to the CPI of most countries. Hence, any ISPI, both intranational and international, will likely to be biased for common reasons. Our numerical experiments suggest that only the path-dependence of CPI can produce bias on a sufficiently large scale as identified here.

This seems to be one more clue to the “PPP puzzle” posed by Rogoff (1966). A failure of time-series testing PPP may be an artifact caused by biases in relative CPIs involved, and not the result of price behavior. Using absolute price data across all EU countries, Crucini, Telmer, and Zachariadis (2005) provide evidence which counts more in favor of PPP than abundant time-series evidence based on the use of CPIs.

At low inflation rates (say, 3% per year), the effect of biases in cross-space comparisons eventually emerges over the long run. As Bessonov (1998) argues, the years of the Russian transition are equivalent to the long run in established market economies due to much faster intrinsic economic time. Thus, the case of Russia can be deemed as a kind of natural experiment that provides insights into what can happen with the ISPI over decades in a developed market economy.

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Appendix A: Results for 12-month averages of incomes

Figure A1. Distributions of 12-month averages of regional incomes (kernel estimates).

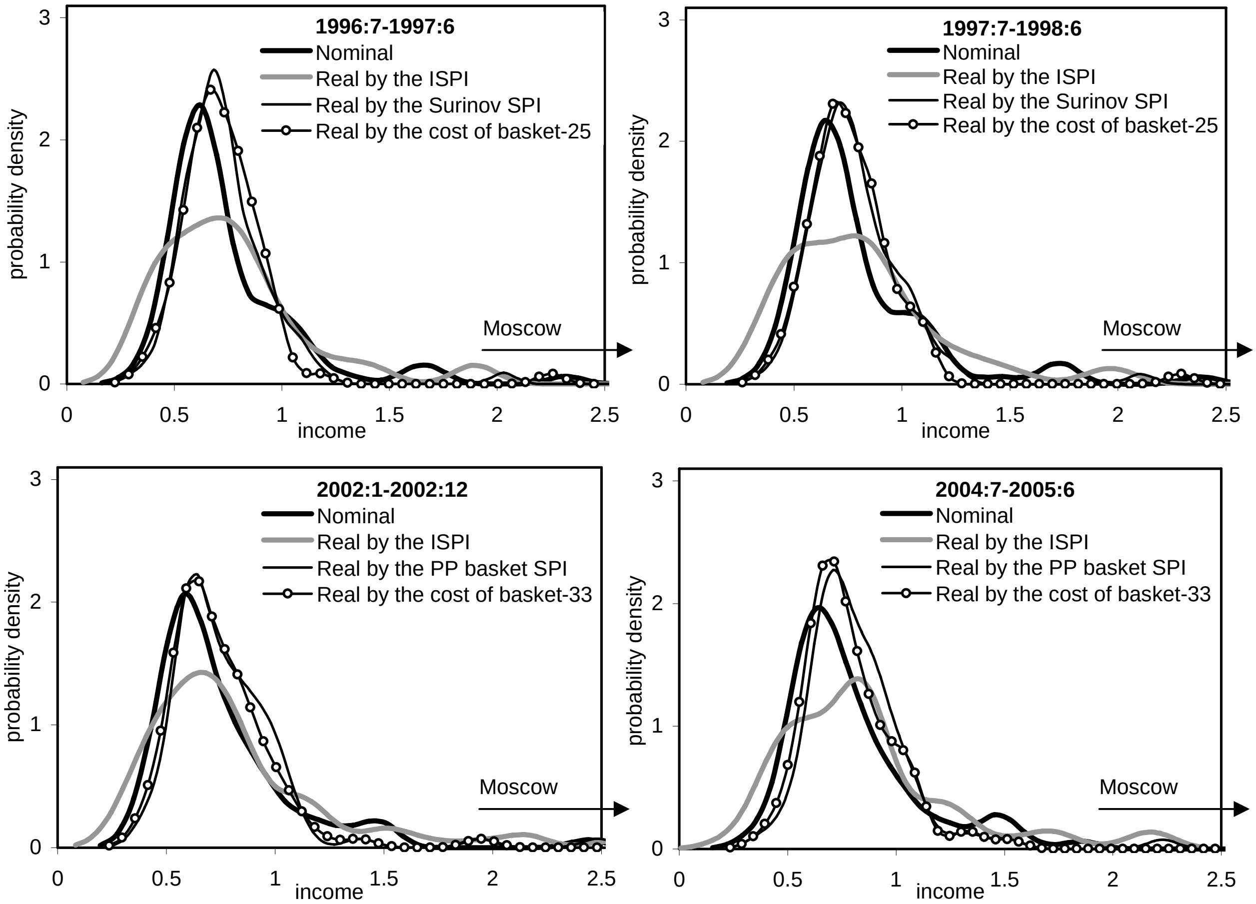

Table A1. Summary statistics for income distribution

Version of real income estimate	Mean	Median	Standard deviation	Minimum	Maximum		
					Without outliers	Tyumen oblast	Moscow
1996:7-1997:6							
Nominal (*M*)	0.793	0.663	0.469	0.323	1.715	2.320	3.739
M/ISPI	0.794	0.709	0.440	0.272	1.909	1.917	3.151
M/DSPI	0.770	0.700	0.377	0.323	1.180	2.030	3.402
M/cost of basket-25	0.765	0.705	0.370	0.349	1.174	2.248	3.217
1997:7-1998:6							
Nominal (*M*)	0.836	0.697	0.469	0.361	1.745	2.383	3.693
M/ISPI	0.838	0.755	0.443	0.300	2.030[a]	1.936	3.046
M/DSPI	0.821	0.753	0.375	0.369	1.230	2.105	3.360
M/cost of basket-25	0.803	0.733	0.358	0.387	1.116	2.287	3.066
2002:1-2002:12							
Nominal (*M*)	0.792	0.644	0.478	0.356	1.514	2.437	3.829
M/ISPI	0.817	0.712	0.442	0.284	2.128	2.168	2.541
M/cost of PP basket	0.781	0.723	0.301	0.398	1.391	1.935	2.481
M/cost of basket-33	0.772	0.687	0.346	0.384	1.390	1.946	2.982
2004:5-2005:6							
Nominal (*M*)	0.830	0.714	0.413	0.318	1.587	1.855	3.341
M/ISPI	0.861	0.804	0.415	0.209	2.194[b]	1.712	2.213
M/cost of PP basket	0.824	0.771	0.254	0.367	1.364	1.514	2.229
M/cost of basket-33	0.807	0.739	0.299	0.374	1.342	1.518	2.705

[a] Vologda Oblast
[b] Republic of Yakutia

Appendix B: Biases of Young index relative to Divisia index

Figure B1. Distributions of Young index biases (kernel estimates).

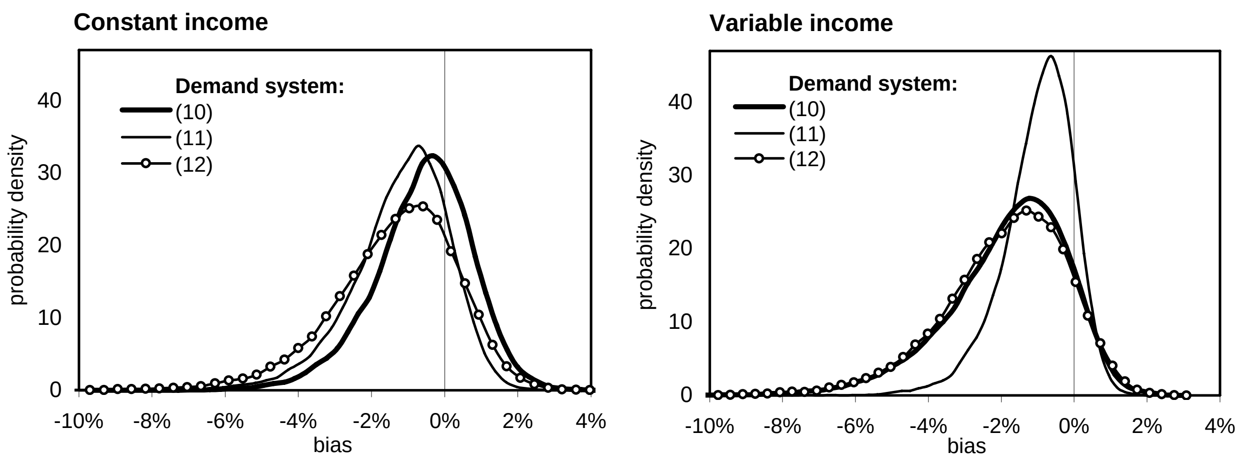

Table B1. Summary statistics for Young index bias distributions (%)

Demand system	Nominal income	Mean	Median	Minimum	Maximum	Standard deviation
(10)	constant	–0.6	–0.5	–9.3	4.3	1.4
	variable	–1.8	–1.6	–13.0	2.8	1.7
(11)	constant	–1.2	–1.0	–10.5	3.5	1.3
	variable	–1.0	–0.9	–8.6	2.1	1.0
(12)	constant	–1.3	–1.2	–13.5	4.3	1.7
	variable	–1.9	–1.7	–13.0	2.6	1.7

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