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Working Paper

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LICOS Discussion Paper, No. 152

Provided in Cooperation with:

LICOS Centre for Institutions and Economic Performance, KU Leuven

Suggested Citation: Gluschenko, Konstantin (2004) : The Law of one Price in the Russian Economy, LICOS Discussion Paper, No. 152, Katholieke Universiteit Leuven, LICOS Centre for Transition Economics, Leuven

This Version is available at:

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LICOS Discussion Papers

Discussion Paper 152/2004

The Law of one Price in the Russian Economy

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The law of one price in the Russian economy^{*}

Konstantin Gluschenko^{}**

Abstact

Taking the law of one price as a test for market integration, the spatial set-up of Russia's market integration over 1994-2000 is analyzed with the use of time series of the cost of a staples basket across Russian regions. The law is found to hold for about 50% to 60% of Russian regions, estimates of a threshold model suggesting rather high barriers to inter-regional trade. To reveal whether there is a movement towards market integration among non-integrated regions, dynamics of cross-sectional distribution of prices receives study. The results indicate that such a tendency does take place. An effort is made to identify forces responsible for inter-regional price disparities.

JEL classification: P22, P25, R15, R19

Keywords: market integration, price dispersion, price convergence, Russia, Russian regions

^{*} This work was supported by an individual grant No. R02-0961 from the Economic Education and Research Consortium, Inc. (EERC), with funds provided by the Eurasia Foundation (with funding from the US Agency for International Development), the Open Society Institute/Soros Foundation, the World Bank, the Global Development Network, the Carnegie Corporation, the Government of Finland, and the Government of Sweden. The author is grateful to participants of the EERC workshops and the International Economics Seminar at the University of Pittsburgh for helpful discussions, and to Michael Beenstock for his comments.

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1. INTRODUCTION

Among stylized facts regarding prices in the transition, Koen and De Masi (1997) argue that, over time, prices have converged across regions within transition countries. Indeed, Gluschenko (2003, 2004a) finds the second half of 1994 to be a turning point from price divergence across Russian regions to the improvement in market integration in Russia. Nevertheless, the Russian market still is not near to be completely integrated. It is apparent that integration in Russia is spatially heterogeneous. For example, difficult-to-access regions markedly contribute to the overall disconnectedness of regional markets. However, the cross-sectional approach applied in the above studies has prevented clarifying the spatial pattern of integration in detail, averaging it over all covered regions of the country.

It is to obtain such a pattern and to make an effort of explaining reasons behind it that this paper aims. The source data for the empirical analysis are time series of the cost of a staples basket across 75 regions of Russia for 1994-2000 with monthly frequency, the cost for Russia as a whole being used as a representative of the national market. The relative cost of the basket is analyzed, which is the ratio of regional cost to the cost for Russia as a whole. (In other words, Russia as a whole is taken as a benchmark.) This implies that integration of regional markets with the entire national market is considered rather than pairwise integration of regions.

The conventional AR(1) model is used to test for the law of one price. This analysis is supplemented with estimating a threshold autoregression model (TAR), obtaining estimates of the effect of impediments to inter-regional trade. Having divided regions to those where the law of one price holds, and those where it does not, the next question is whether the latter regions do move towards integration. One more question is whether the Russian market is fragmented to isolated clusters of regions tending to integrate with each other (price convergence clubs). To answer these questions, the cross-sectional price distribution is analyzed. First, dynamics of price dispersion is tested for σ -convergence. Second, the distribution itself is non-parametrically estimated, tracing changes in its shape over time. Third, a transition probability matrix is estimated, which characterizes intra-distribution mobility. Such a matrix is used also to derive a long-run limit of the distribution. At last, forces impeding regional prices to equalize are identified.

The law of one price is found to hold for about 50% to 60% of Russian regions over 1994-2000. The TAR estimates suggest rather high barriers to inter-regional trade that equal, on

average, 13% to 15% of the price in Russia as a whole. Thus, the degree of market integration is not high. However, prices in non-integrated regions do exhibit σ -convergence. The shape of the cross-sectional distribution of prices in those regions tends to be more regular and narrower over time, however, keeping a long right-hand tail that is due, in the main, to difficult-to-access regions. Both this distribution and that across all the 75 regions are unimodal, suggesting the absence of price convergence clubs. The estimated transition probability matrix evidences, too, price convergence during 1994-2000. The long-run limit of the price distribution is unimodal, so suggesting no price convergence clubs in the long run either.

In fact, this work has preceded the empirical analysis described in Gluschenko (2004b). In the latter paper, more sophisticated econometric tools are used that take account of both the specificity of prices in the transition and features of the Russian economy. Nonetheless, the results reported in this paper still are of interest, presenting a pattern yielded by (more or less) traditional methodologies of analysis that are applied to markets in advanced market economies.

Analyzing the law of one price (and market integration in general) in Russia and other transition economies has been the subject of a number of studies.¹ Using time-series analysis, Gardner and Brooks (1994), Goodwin *et al.* (1999), and Berkowitz *et al.* (1998) examine price dispersion among Russian cities in the early years of the transition (up to 1995). They find the Russian market weakly integrated, yet encouraging signs are found as well that a number of integrated city pairs do exist. (An early version of the paper by Berkowitz *et al.* (1998) was even titled “Transition in Russia: It’s Happening.”) Gluschenko (2001) analyzes cointegration and threshold relationships across 7 regions of Western Siberia with the use of the food CPI and industrial goods CPI. A similar analysis is performed by Gluschenko (2002) across all the 11 aggregated economic territories (*ekonomicheskii rayon*) of Russia.² More recently, Gluschenko and Koneva (2004) use the cost of a staples basket instead, comparing results with those yielded by the CPI as a price representative. Both integrated and non-integrated region/territory pairs are found with these analyses, the CPI goods components being revealed to be an inadequate price indicator for analyzing the law of one price.

Cross-sectional approach also finds use in the relevant studies. Exploiting it, Berkowitz and DeJong (1999) reveal a culprit behind the fragmentation of the Russian market, the so-called

¹ Contrary to the opinion of Fan and Wei (2003), and Horvath and Vidovic (2004), this number is not too small.

² The traditional economic zoning of Russia is meant, as it appeared, e.g., in Goskomstat’s publications until July 2000, except for the Kaliningrad Oblast which is added to the Northwestern Territory. This zoning is also referred to in Section 4.1 of this paper.

“Red Belt”; and then Berkowitz and DeJong (2001, 2003) obtain a segment of the integration trajectory for Russia. Gluschenko (2003) also derives such a trajectory for 1992-2000 with the use of a different methodology; Gluschenko (2004a) reports some other estimates of the degree of Russia’s market integration.

Conway (1999), using data from 1993–1996, examines price convergence among four market locations within Kiev, Ukraine. He finds significant evidence of price convergence due to arbitrage by buyers and sellers at these markets, but sizeable and sustained divergences from the law of one price have remained as well. Cushman *et al.* (2001) also examine the law of one price in Kiev, comparing Ukrainian and US prices during 1991–1992. Although the law did not hold, the commodity real exchange rates are found to have possessed deterministic trends that were in the direction of closing the initial price gap. Studying price dispersion in China, Young (2000) finds it to substantially fluctuate over time. Nevertheless, a panel unit root analysis performed by Fan and Wei (2003) suggests that the law of one price does hold in China, the pattern being highly comparable with that in the US and Canada. Ratfai (2003) also performs unit root tests over a number of product panels, each covering 120 months and 20 locations in Hungary. He finds the law of one price to hold for each product; the implied half-lives in general show very fast (short-run) convergence to the price parity. Horvath and Vidovic (2004) examine prices across districts of Slovakia over a five-year period. They find evidence for convergence to the law of one price; however, its speed is lower than that found for US cities.

Worthy of mention are also papers by Vincent Koen with co-authors, in particular, De Broeck, De Masi and Koen (1995), De Masi and Koen (1996), and Koen and De Masi (1997). They touch on the issue of geographical price dispersion in transition economies as well, although without the use of econometrics.

This paper also relates to papers analyzing domestic market integration in advanced market economies, such as Engel and Rogers (1996), Parsley and Wei (1996), Obstfeld and Taylor (1997), and Rogers (2002). More distantly, it relates as well to countless papers on analysis of the law of one price in the international context and purchasing power parity (PPP), most sufficient of which were surveyed by Rogoff (1996) and Sarno and Taylor (2002).

2. METHODOLOGY

2.1. Strategy of the analysis

Perfect integration of a geographically dispersed goods market implies that there are no impediments to the movement of goods between all its spatial segments, e.g., regions of the country. In other words, perfectly integrated market operates like a single market despite its spatial dispersion. Then the price of a (tradable) good across regions is uniform, i.e., the law of one price holds, inter-regional arbitrage maintaining the law to hold. Thus, the law of one price can be used as a theoretical benchmark for empirically analyzing domestic market integration.³

As mentioned in Introduction, there are two stages in the evolution of market integration in Russia, namely, the early stage of progressive disintegration beginning in January 1992, and the late stage of improvement in integration beginning in about 1994. It is the late stage that is of interest in this study. It is hypothesized that the Russian goods market should eventually come to the final steady state of complete integration, that is, to the equality of prices across all regions. Currently, during the late stage of the integration evolution, the market may be believed to be in transition towards this steady state.

Hence, it is expected that three groups of regions can currently exist: (1) integrated regions, i.e., those being in the steady state of equality of prices; (2) non-integrated regions tending towards integration, i.e., those in which prices are catching up with each other; and, maybe, (3) non-integrated regions having no such a trend. Along with this, either of groups (1) and (2) could break into subgroups of mutually integrated (respectively, tending to mutual

³ While the law of one price as the theoretical benchmark for integration, and arbitrage transaction costs as the culprit behind violations of the law are commonly used in the international and intra-national context, some economists are skeptical of this, basing on patterns of within-city price dispersion. Theoretical models, e.g., by Salop and Stiglitz (1977) and Varian (1980), to name a few, predict within-city price dispersion to be permanent, transaction costs not being a culprit (at least, the only one); Lach's (2002) empirical results confirm the Varian model. However, inter-location and within-city price dispersions are rather different stories. There are many effects possible between stores within a city but impossible between, say, cities of a country. For example, a uniform good can be differentiated by transforming it into a combination "the good + the prestige of a store". Varian's (1980) model provides another example: each store varying its price over time (switching from lower to higher price and *vice versa*) in order to disorient consumers. There is no free entry to many (if not most of) spatial segments of the within-city market; hence, arbitrage opportunities can remain unrealized for arbitrage itself is hardly possible in such cases. On a higher level, e.g., that of the national market, where entire city or region is treated as a single market agent, the structure of the market becomes much simpler. It is believed that all barriers to trade in such a market can be interpreted in terms of arbitrage transaction costs in the broad sense (of course, unless the case in hand is a monopolistic market for some good). For instance, such a barrier like local trade restrictions (which were introduced by regional authorities in some Russian regions, and Young (2000) reports that the same took place in China) can be quantified as costs needed for cutting through the bureaucratic red tape.

integration) regions; such subgroups are an analog of convergence clubs in economic growth; see, e.g., Barro and Sala-i-Martin (1995). This implies that the steady state of the entire market would split into multiple steady states for region clusters, thus preventing the market to be completely integrated even in its steady state (in fact, the market will be fragmented to a few isolated submarkets which are internally integrated). These are possible features of the pattern of Russia's market integration. Revealing the actual pattern is the main objective of this paper.

In the above context, the term *convergence of prices* becomes ambiguous. In fact, considering region groups (1) and (2), two fundamentally different concepts of convergence come into collision with one another. Let P_{rt} (P_{st}) denote log price of a good in region r (s) at time $t \in [0, T]$, and $P_{rst} \equiv P_{rt} - P_{st}$ denote (percent) price differential. Figure 1 illustrates the difference between the concepts; thin lines in the figure depict actual dynamics of prices; and thick lines represent their theoretical (long-run) trajectories.

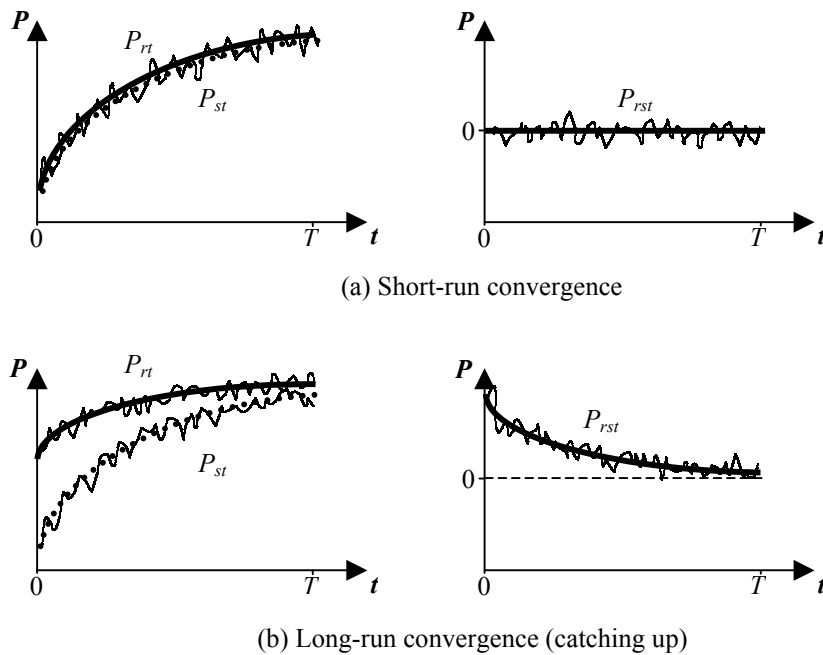


Figure 1. Two concepts of price convergence

In words, these two concepts can be described as follows:

(a) Figure 1(a) implies that regions r and s fall into group (1) of those *being* in the spatial equilibrium; price disparities between regions are merely random shocks dying out over time. In other words, prices fluctuate around parity, permanently tending to return to it. This is just the case that the literature on the law of one price and PPP deals with, calling it “convergence to the

law of one price/PPP.” That is, the term “convergence” relates to shocks, implying their convergence to zero. In fact, this characterizes the short-run behavior of prices, while their long-run behavior is described by the trajectory $P_{rst} = 0$. Thus, such a concept can be called “short-run convergence.”

(b) Figure 1(b) implies that regions r and s fall into group (2) of those *tending* towards the spatial equilibrium: $\lim_{t \rightarrow \infty} P_{rst} = 0$; in the figure, prices in s are catching prices in r up. In general, the price disparity permanently diminishes over time; and the prices fluctuate around this general trend due to random shocks. This is the case that the literature on economic growth deals with (regarding income data), calling it simply “convergence.” This time, in the short run, prices converge to the long-run trajectory (i.e., random deviations from it dye out over time), while this trajectory itself converges to the parity line $P_{rst} = 0$ in the long run. It is the latter tendency that is of main interest. Here, the term “convergence” relates to prices themselves, implying (long-run) convergence of their differences to zero over time. Hence, such a concept can be called “long-run convergence.”

Analyzing short-run price convergence, that is, testing for the law of one price, is the standard exercise in the voluminous literature on this issue. In essence, it is testing time series $\{P_{rt}\}_{t=0, \dots, T}$ and $\{P_{st}\}_{t=0, \dots, T}$ for cointegration with the prespecified cointegrating vector $(1, -1)$. Because of disturbances, $P_{rst} = v_{rst}$ will be observed rather than exact parity (where v_{rst} is a stochastic variable), which leads the problem to testing P_{rst} for stationarity. That is, the question is whether v_{rst} is really the equilibrium error. Assuming v_{rst} to be a (first-order) autoregressive process, $v_{rst} = (\lambda_{rs} + 1)v_{rs,t-1} + \varepsilon_{rst}$, a testable version of the theoretical relationship $P_{rst} = 0$ is arrived at (with $\Delta P_{rst} = P_{rst} - P_{rs,t-1}$):

$$\Delta P_{rst} = \lambda_{rs} P_{rs,t-1} + \varepsilon_{rst} \quad (\text{across } t = 1, \dots, T \text{ with fixed } (r, s)). \quad (1)$$

If $-2 < \lambda_{rs} < 0$ then the law of one price holds, hence, regions r and s are integrated.

Analyzing long-run price convergence, or trend towards integration, is much more challenging. The expression $\lim_{t \rightarrow \infty} P_{rst} = 0$ means that process $\{P_{rst}\}_{t=0, \dots, T}$ is not stationary, but tends to a stationary one over time. The literature does not provide any statistical test for such a kind of processes with the use of a univariate time-series model. (Except for the case when $\{P_{rst}\}$ becomes stationary since some $\tau \in (0, T]$. Then a test could be a sequence of regressions (1) across $t \in [0, T]$ with θ varying from 1 to τ .) Therefore, this should be done cross-sectionally. It is

a traditional problem for analysis of economic growth, but as far as I know, it has not so far found use in analysis of market integration.⁴ (A possible reason may be that, usually, the case of interest is integration as an established state, while dealing with a transition economy, the movement from non-integration to integration is to be expected, and holds the greatest interest.)

Two main notions of long-run convergence appear in the economic growth literature, termed β -convergence and σ -convergence (Barro and Sala-i-Martin, 1995).

Reformulating in terms of prices, if the regression⁵

$$\Delta P_{rst} = \beta P_{rs,t-1} + \varepsilon_{rst} \quad (\text{across all region pairs } (r, s) \in \Pi; t \text{ and } t-1 \text{ are fixed}) \quad (2)$$

yields $\beta < 0$ then it is said that the data set exhibits β -convergence.⁶ Set Π is that of unique region pairs, i.e., if $(r, s) \in \Pi$ then $(s, r) \notin \Pi$, and $r \neq s$. For example, this set can be constructed as $\Pi = \{(r, s) \mid r > s\}$. Time points t and $t-1$ in (2) are not assumed to be neighboring ones; they can be rather distant, being separated by a long time interval.

However, as shown by Friedman (1992), Quah (1993b), Bernard and Durlauf (1996), Wodon and Yitzhaki (2001), etc., β -convergence is not informative for actual behavior of the cross section, suffering from Galton's fallacy. Though being consistent with the pattern in Figure 1(b), $\beta < 0$ can be consistent as well even with diverging prices.⁷

Another notion, σ -convergence, is defined as follows (again, reformulated in terms of prices): regional prices are catching up with each other if their dispersion tends to decrease over time, that is,

$$\sigma(P_t)/\sigma(P_{t-1}) < 1 \quad (\text{with fixed } t \text{ and } t-1), \quad (3)$$

where $\sigma(P_t)$ is the standard deviation of the price differential over all $(r, s) \in \Pi$ for a given point

⁴ Michael Beenstock brought this idea to my attention.

⁵ In the economic growth literature, this regression uses locations' levels rather than inter-location disparities, testing the hypothesis that the lower initial (at $t-1$) level, the faster growth. In terms of disparities, like in (2), this is equivalent to that the more initial disparity, the faster decrease of the gap.

⁶ In fact, though not mentioned in the growth literature, $\beta > -2$ should hold as well.

⁷ The similarity of Equations (1) and (2) to one another might cause a suspicion that Equation (1) can also lead to Galton's fallacy. But this is not the case: it is the difference in the data sets over which (1) and (2) are run that is fundamental. In regression (2), observations in only two points in time are analyzed, while Galton's fallacy demonstrates that it is inappropriate to draw dynamic implications from cross-section evidence. With (1), the story is quite different, as just the dynamics is analyzed over a sequence of time points, thus taking into account ongoing disturbances (the lack of which just gives rise to Galton's paradox). It is easily shown formally, applying Quah's (1993b) proof (that the sign of β in Equation (2) says nothing about whether there is convergence or divergence) to Equation (1), that the negative sign of λ_{rs} is consistent only with convergence of deviations from the law of one price to zero, non-negative sign saying about divergence (randomly walking or permanently rising $\{P_{rst}\}$). See also Bernard and Durlauf (1996) for further discussion of the difference between results yielding by (1) and (2).

in time (the aforesaid as to the time points $t-1$ and t holds). Unlike β -convergence, σ -convergence reveals actual properties of the behavior of prices. Therefore examining σ -convergence is adopted in this paper, being applied to both the entire region sample and to the subsample of regions found to be non-integrated.

However, analyzing σ -convergence does not answer the question of whether there are price convergence clubs. To reveal more detailed properties of the integration evolution, the behavior of the cross-section distribution of regional prices as such, $f_t(P)$, receives study. Two ways are used. The first is non-parametrically estimating distribution $f_t(P)$ for a number of $t \in [0, T]$ in isolation. The second way is estimating transition probability matrix $\mathbf{M}(\tau)$, following Quah (1993a). For this, the range of price differences is discretized by dividing it into a number of “price classes.” An element $m_{ij}(\tau)$ of the matrix is the probability of regions/region pairs to transit to price class j (with price differentials from $P_1^{(j)}$ to $P_2^{(j)}$) at time $t + \tau$, being in price class i at t . (Thus, $\mathbf{M}(\tau)$ contains information on intra-distribution mobility.) Having such a matrix, a long-run limit of the price distribution can be estimated. In distribution terms, the issue of convergence clubs is that of unimodality vs. multimodality of a distribution.

Unfortunately, the distribution dynamics analysis gives no way to clearly classify non-integrated regions into groups (2) and (3), i.e., into the group of regions tending to integration and those having no such a trend. Results of the analysis can suggest only whether the trend to convergence of prices prevails over the trend to divergence induced by regions from group (3), or whether it does not.⁸

This paper deals with 75 regions of Russia, which would yield 2775 region pairs. To reduce such a mass of pairwise comparisons, the national market as a whole is taken as a benchmark in the empirical analysis. Thus, integration of each region with the entire national market is analyzed, using only region-Russia pairs rather than all region-by-region pairs. (The latter is reserved for future research.) That is, index s in the above relationships is, in fact, fixed, and is set to $s = 0$, P_{0t} denoting the price in Russia as a whole. (Then $P_{rst} \equiv P_{rt} - P_{0t}$ is the deviation of the regional price from the national average, or the regional price relative the national one.)

Having outlined the framework of the analysis in this section, Sections 2.2 through 2.4 describe the methodology in more detail. Testing for the law of one price is supplemented with

⁸ Subsequently, Gluschenko (2004b) has found a way of solving this problem.

additional time-series analyses presented in Section 2.2. Section 2.3 is concerned with the distribution dynamics analysis. At last, estimates yielded by the time-series analyses are used as endogenous variables in cross-section regressions aiming to explain the pattern obtained. Section 2.4 describes this analysis, arbitrarily called the “explanatory analysis.”

2.2. Time-series analysis

The conventional autoregression model AR(1) specified by (1) that is used to test for the law of one price, is estimated individually for each region r ; $r = 1, \dots, R$. The Phillips-Perron test is applied to test time series for unit root, 3 lags being taken (the Newey-West truncation lag for a 83-observation time series). This test is chosen in order not to lose the degrees of freedom by inserting additional lags into the equation itself. The estimate of λ_{rs} is interpreted as the short-run convergence speed, that is, the speed of dying out deviations from the law of one price. A more descriptive derivative indicator is calculated as well, the half-life time of the deviations:

$$\theta_{rs} \equiv \theta(\lambda_{rs}) = \ln(0.5)/\ln(|1 + \lambda_{rs}|). \quad (4)$$

The strict law of one price, $P_{rst} = 0$, is too rigid, assuming arbitrage to drive inter-regional price differences to exact zero. It is more realistic to assume that there are arbitrage transaction costs preventing such an exact equalizing. When the price difference becomes less than these costs, arbitrage turns out unprofitable, and so, does not act. Thus, the weak version of the law of one price is arrived at: $C_{(-)rs} \leq P_{rst} \leq C_{(+)rs}$, where $C_{(-)rs} = \ln(1 - c_{rs})$, $C_{(+)rs} = \ln(1 + c_{rs})$, and c_{rs} is arbitrage transaction costs expressed as a percentage of the price. (Note that $C_{(-)rs} < 0$ and $C_{(+)rs} > 0$.) The weak law allows for persistent price dispersion varying within the band $(C_{(-)rs}, C_{(+)rs})$, the entire band being the spatial equilibrium. The asymmetry of the thresholds is not a usual practice; Appendix A provides theoretical considerations behind it.

The threshold autoregression model (TAR) is adopted as a testable version for the weak law of one price. Assuming dynamics of the price differential to combine two components, a wittingly stationary processes outside the band $[C_{(-)rs}, C_{(+)rs}]$, and an arbitrary (maybe, non-stationary) process within the band, the TAR explicitly pulls these components apart:

$$\Delta P_{rst} = \begin{cases} \lambda_{rs}^* (P_{rs,t-1} - C_{(+rs)}) + \varepsilon_{rst}^* & \text{if } P_{rs,t-1} > C_{(+rs)} \\ \lambda_{rs}^0 P_{rs,t-1} + \varepsilon_{rst}^0 & \text{if } C_{(+rs)} \geq P_{rs,t-1} \geq C_{(-rs)} \\ \lambda_{rs}^* (P_{rs,t-1} - C_{(-rs)}) + \varepsilon_{rst}^* & \text{if } P_{rs,t-1} < C_{(-rs)} \end{cases} \quad (5)$$

$$-2 < \lambda_{rs}^* < 0.$$

The model is estimated individually for each region r ($r = 1, \dots, R$) across $t = 1, \dots, T$.

To estimate and test the TAR model, a maximum likelihood approach is applied, somewhat modifying the method put forward by Obstfeld and Taylor (1997). The essence of it is a greed search of the value of the threshold with the use of the log ratio of the likelihood function of the TAR model and that of the AR model, LLR, as a criterion. Following Andrews (1998), allowable values of the threshold are set so that sets $\{P_{rst}\} \in [C_{(-rs)}, C_{(+rs)}]$ and $\{P_{rst}\} \notin [C_{(-rs)}, C_{(+rs)}]$ contain not less than 15% of observations. The test for the threshold effect is, in fact, a specification test of (5) against (1).⁹ LLR is used as the test statistic as well. To derive its p -value, a model-based bootstrap is applied with 1,000 replications. Hereafter this LLR-based test is referred to as the LLR test.

The results are an estimate of λ_{rs}^* , the speed of dying out deviations from the equilibrium, and an estimate of threshold, $C_{(+rs)}$ or $C_{(-rs)}$ (which is the same as they are functionally dependent; see Appendix A). The half-life time for the TAR is calculated by Formula (4) with the use of λ_{rs}^* instead of λ_{rs} : $\theta_{rs}^* = \theta(\lambda_{rs}^*)$. For the ease of interpretation, the estimated thresholds are transformed to percentage terms as

$$c_{rs} \equiv c(C_{(-rs)}) = 1 - \exp(C_{(-rs)}) = \exp(C_{(+rs)}) - 1. \quad (6)$$

The threshold represents the total effect of all impediments to trade between r and s . But, accordingly to the definition in the beginning of Section 2.1, it is the absence of such impediments that is inherent in the integrated market. Hence, in the general case, the non-rejection of the threshold effect does not indicate integration between r and s . And so, the main purpose of this way of the analysis is just estimation of the total effect of forces preventing regions from integration.

There is a caveat with this. Of course, perfect integration is not observed in the real world (at least, this is not a common instance): e.g., Engel and Rogers (1996) as well as Parsley and

⁹ Since model (5) implies that deviations from the (weak) law of one price converge to external borders of the arbitrage inaction band $[C_{(-rs)}, C_{(+rs)}]$ rather than to zero, if there is no stationary process outside this band then the TAR specification is rejected in favor of AR(1) with a unit root.

Wei (1996) find price dispersion among US cities to depend strongly on distance. Therefore, to be more realistic, the criterion for market integration could be weakened, allowing for “natural”, irremovable impediments to inter-regional trade such as physical distance and difficult access to a number of regions. But this requires a knowledge of reasons behind arbitrage specific transaction costs: when $\lambda_{rs} = 0$ and the threshold effect is not rejected, r and s could be deemed as integrated only if c_{rs} is nothing but the cost of shipping the good between r and s (and/or the effect of difficult access to the region). Unfortunately, there is no way to ascertain this for individual regions (as well as region pairs).

If regression (1) were run for all region pairs $(r, s) \in \Pi$, this would make feasible revealing price convergence clubs by an analysis of *indirect integration*. As s is fixed, such a method is not exploited in this paper; it is described for completeness sake.¹⁰

The set of results for all region pairs indicate what regions are integrated with one another, and what are not. This pattern can be described by an $R \times R$ “integration matrix” $\mathbf{J} = (J_{rs})$ such that $J_{rs} = 1$ if $\lambda_{rs} < 0$, and $J_{rs} = 0$ if $\lambda_{rs} = 0$; $J_{rr} \equiv 1$. In fact, this is the adjacency matrix of a graph displaying direct price linkages between regions. However, indirect linkages can exist as well. As Goodwin *et al.* (1999) note, overlapping regional markets may maintain spatial linkages even in instances when direct trade does not occur between a pair of regions. In other words, if regions r and s are not integrated with one another in the sense of the law of price test, but the both are integrated with the same third region q , it is reasonable to conclude that integration does take place between r and s (as prices in r and s are fluctuating about some common level), but it is low power of a test for statistical significance of the λ_{rs} (unit root test) that prevents detection of this due to moderate magnitude of λ_{rs} . Hence, these regions r and s can be deemed as “indirectly integrated” through region q .

This means that there is a chain connecting r and s in the “integration graph” represented by matrix \mathbf{J} . In this case, the chain length is 2; certainly, longer chains may connect “seemingly non-integrated” regions. The n -th power of \mathbf{J} provides a pattern of indirect (or “chained”) integration of “the n -th degree”; its element $J^{(n)}_{rs}$ is nonzero if there exists a chain of length no more than n between r to s , otherwise $J^{(n)}_{rs} = 0$. With $n = R-1$, a clear-cut pattern of the “indirect integration” is yielded. A zero rs -th element of \mathbf{J}^{R-1} implies that there are neither direct nor indirect price linkages between r and s .

¹⁰ This method is applied by Gluschenko (2002) and Gluschenko and Koneva (2004).

This pattern answers to the question whether the integrated part of the national goods market is single or whether it is fragmented to a few isolated submarkets. Let regions be reordered so that \mathbf{J}^{R-1} becomes a block diagonal matrix; if there is more than one block containing non-zero non-diagonal elements, then the latter is the case. Isolated submarkets represented by such blocks are just convergence clubs.

2.3. Distribution dynamics analysis

The first issue of interest is σ -convergence. As seen from Formula (3), testing for it is rather trivial. Relatively distant points in time, e.g., those separated by 12-months intervals, can be compared to check whether (3) is valid. For more thorough analysis of σ -convergence, $\{\sigma_t\}_{t=0,\dots,T}$ can be treated as a time series, testing it for unit root. If the series proves to be trend stationary with a negative trend factor, this suggests σ -convergence.

Then the degree to which the shape of the distribution of the price differential changed over time is traced. To assess these changes, the cross-sectional distributions are non-parametrically estimated in a number of points in time with the use of a kernel density estimator. The Gaussian kernel is adopted; formally, the estimate of a probability density looks like

$$\hat{f}_{St}(P) = \frac{1}{R_S h \sqrt{2\pi}} \sum_{r \in \Pi_S} \exp\left(-\frac{1}{2} \left(\frac{P - P_{r0t}}{h}\right)^2\right), \quad (7)$$

where S indexes region samples under consideration, Π_S is the set of regions (more exactly, of “region-Russia” pairs) in S -th sample, R_S is the number of regions in this sample, and h is a smoothing bandwidth.

Having obtained a sequence of the distributions, the transition process between them, i.e., price mobility of regions, receives study. In brief, the method of this analysis is as follows (see Quah (1993a) and Durlauf and Quah (1999) for its substantiation and technical description). Let $f_t(P)$ be the distribution of prices (related to the Russian average) at time t . Its evolution can be described as

$$f_{t+\tau}(P) = M(\tau) f_t(P), \quad (8)$$

where $M(\tau)$ is an operator mapping one distribution to another. The operator is assumed to be time-invariant; then, applying transformation (8) to $f_t(P)$ k times, a distribution for $t+k\tau$ is obtained, i.e.,

$$f_{t+k\tau}(P) = M^k(\tau) \cdot f_t(P). \quad (9)$$

Taking $k \rightarrow \infty$ yields the ergodic distribution, $f_\infty(P)$, that is, such that

$$f_\infty(P) = M_\infty \cdot f_\infty(P), \quad (10)$$

where M_∞ is the limit of $M^k(\tau)$ with $k \rightarrow \infty$. The ergodic distribution is a long-run limit of the distribution of prices across regions.

Let the price range be divided into n cells (price classes), so providing the discretization of distribution $f_t(P)$: it transforms to a (raw) vector $\mathbf{f}_t = (f_{it})$, where $i = 1, \dots, n$ is the class number; i -th element of the vector is the probability of the price falling into class i . Operator $M(\tau)$ changes into an $n \times n$ Markov chain transition matrix $\mathbf{M}(\tau) = (m_{ij}(\tau))$ whose ij -th element is the probability that a price from class i transits to class j over time τ . Being sequentially powered, transition probability matrix $\mathbf{M}(\tau)$ will be converging to a rank-one matrix \mathbf{M}_∞ having equal rows; these rows are just equal to the ergodic probability vector, \mathbf{f}_∞ .¹¹

2.4. Explanatory analysis

The above strands of the analysis yield a pattern describing the state and tendencies of Russia's market integration, identifying what part of the Russian market is already integrated, and whether the rest part of the market is moving towards integration. However, they do not provide an explanation for the pattern. And so, this part of the analysis could be arbitrarily labeled as “descriptive.” The next part, equally arbitrarily labeled as “explanatory,” addresses just this issue: what it is that explains the features of the pattern obtained. To put it differently, it aims to reveal forces responsible for the features found.

The variables to be explained are estimates yielded by the “descriptive part” of the analysis, namely, by the time-series analysis. Variables that are potentially explanatory (and for which information is available) are listed in Section 4.3; in formulas below, they are denoted X_{rsk} , where k indexes specific variables. Coefficients on them are denoted α_k (implying α_0 to be the intercept, and $X_{rs0} \equiv 1$).

To estimate roles played by various forces in the fact that regions are not integrated, a binary choice (logit/probit) model is adopted,

¹¹ Durlauf and Quah (1999) provide a proof of this.

$$\Pr(J_{rs} = 0) = \Psi\left(\sum_k \alpha_k X_{rsk}\right), \quad (11)$$

where $\Psi(\cdot)$ denotes a cumulative distribution function; event J_{rs} represents a result of testing for the law of one price with the use of Equation (1): $J_{rs} = 0$ if $\lambda_{rs} = 0$ and $J_{rs} = 1$ if $\lambda_{rs} < 0$. (In fact, J_{rs} is an element of integration matrix \mathbf{J} .) A modification is also tried, based on the fact that the variable determining the choice between $J_{rs} = 0$ or $J_{rs} = 1$ is not actually latent, since $J_{rs} = 0$ if $p(\lambda_{rs} = 0) > P_L$, where $p(\cdot)$ is the unit root test p -value, and P_L is a critical significance level (namely, 10%). Thus, the p -value can be directly used as the dependent variable, and the ordinary linear model is arrived at,

$$p(\lambda_{rs} = 0) = \sum_k \alpha_k X_{rsk} + \varepsilon_{rs}. \quad (11')$$

The purpose of analyzing barriers to trade is to explain the total effect of impediments to inter-regional trade, i.e., to reveal contributions of various forces to thresholds c_{rs} estimated with the use of TAR model (5). A linear model is adopted for this,

$$\hat{c}_{rs} = \sum_k \alpha_k X_{rsk} + \varepsilon_{rs}. \quad (12)$$

If the variables representing transportation costs and the difficult access to some regions only are significant, the market as a whole can be deemed as integrated.

Turning to indirect integration, the object of interest could be the degree of integration (the length of the shortest chain between two regions in the integration graph). Let it be denoted D_{rs} ; $D_{rs} = d$ if $J_{rs}^{(d)} > 0$, but $J_{rs}^{(d-1)} = 0$; $d = 1, \dots, D^*$. (Usually, the maximum degree of indirect integration, D^* , is less than $R-1$.) For convenience, D^*+1 means the absence of indirect integration of any degree. To explain the pattern of indirect integration, the multiple (ordered) choice model can be adopted,

$$\Pr(D_{rs} = d) = \Psi\left(\sum_k \alpha_k X_{rsk}, d\right) \quad (d = 1, \dots, D^* + 1). \quad (13)$$

However, since this paper deals only with pairs $\{(r, 0)\}$, this model is not actually used in it.

3. DATA

The cost of the basket of 25 food goods (defined as the standard by Goskomstat between January 1997 through June 2000) is used as a price representative for the analysis. This basket covers

about one third of foodstuffs involved in the Russian CPI; but unlike the CPI, it has constant weights across regions and time. Goskomstat (1996a) describes its composition. The costs of the basket were obtained directly from Goskomstat's office (the index before 1997 was computed there at the request of the author; as for July to December 2000, Goskomstat still continued calculating this index though did no longer publish it).

The data are collected in capital cities of the Russian regions; 75 of 89 Russia's regions are covered ($R = 75$). The data are lacking for 10 autonomous *okrugs*, the Chechen Republic, and the Republic of Ingushetia. Besides that, two more regions are omitted. The city of Moscow is a "city-region", being a separate subject of the Russian Federation, and at the same time it is the capital city of the surrounding Moscow Oblast; the same holds for St. Petersburg and the Leningrad Oblast. Therefore only these "cities-regions" are present in the sample, while the relevant surrounding *oblasts* are not. The data are monthly, spanning January 1994 through December 2000. Thus, the number of the time observations equals 84 ($T = 83$, beginning from $t=0$).

There are missing observations in the time series used. To fill the gaps, missing prices are approximated, using the food component of the regional monthly CPIs. The interpolated value of price $p_r(t)$ is the arithmetic mean of the nearest known preceding price inflated to the required time point, t , and the nearest known succeeding price deflated to t . Let prices at $t - m$ and $t + n$ be known, and $\pi_r(\tau)$ be CPI-food for month τ in region r ($\pi_r(\tau) = \bar{p}_r(\tau) / \bar{p}_r(\tau - 1)$, where $\bar{p}_r(\cdot)$ is the overall level of prices for foodstuffs). Then the interpolated price is computed as

$$p_r(t) = \frac{1}{2} (p_r(t - m) \cdot \pi_r(t - m + 1) \cdot \dots \cdot \pi_r(t) + \frac{p_r(t + n)}{\pi_r(t + n) \cdot \pi_r(t + n - 1) \cdot \dots \cdot \pi_r(t + 1)}). \quad (14)$$

For example, if an observation for one month is missed, its restored value looks like $p_r(t) = (p_r(t-1) \cdot \pi_r(t) + p_r(t+1) / \pi_r(t+1)) / 2$.

Below, regions with difficult access are referred to. These regions are: the Murmansk Oblast, the Republic of Sakha (Yakutia), the Sakhalin Oblast, the Magadan Oblast, and the Kamchatka Oblast. They are remote regions lacking (except the Murmansk Oblast) railway and highway communication with other regions. Due to this, arbitrage can hardly be bilateral there, goods being imported only to these regions.

4. EMPIRICAL RESULTS

4.1. Testing for the law of one price

Table 1 summarizes results on integration of each individual region with the entire national market, which are obtained with the use of both AR and TAR models. In this table, standard deviations are in parentheses, $p(\lambda=0)$ and $p(\text{AR})$ are p -values of the Phillips-Perron test and the LLR test, respectively. The 10% significance level is adopted; p -values exceeding it are marked with bold italics. Regions are arranged in descending $|\lambda_{r0}|$.

Table 1. Testing for the strict and weak versions of the law of one price

Region	Autoregression model (1)			Threshold model (5)			
	$p(\lambda=0)$	λ	θ , months	$p(\text{AR})$	λ^*	θ^* , months	c , %
Chelyabinsk Obl.	0.000	-0.698 (0.105)	0.6	0.000	-1.902 (0.402)	6.7	6.1
Novgorod Obl.	0.000	-0.469 (0.092)	1.1	0.040	-0.637 (0.140)	0.7	1.8
Rep. of Altai	0.000	-0.401 (0.088)	1.4	0.003	-1.430 (0.418)	0.8	8.5
Samara Obl.	0.000	-0.376 (0.087)	1.5	0.332	-0.825 (0.146)	0.4	4.7
Buryat Rep.	0.001	-0.276 (0.075)	2.1	0.000	-0.798 (0.238)	0.4	10.8
Tomsk Obl.	0.001	-0.252 (0.071)	2.4	0.062	-0.675 (0.145)	0.6	6.1
Kemerovo Obl.	0.000	-0.237 (0.058)	2.6	0.161	-0.418 (0.091)	1.3	3.6
Novosibirsk Obl.	0.000	-0.234 (0.059)	2.6	0.078	-0.509 (0.289)	1.0	7.3
Perm Obl.	0.006	-0.204 (0.067)	3.0	0.001	-1.200 (0.179)	0.4	5.7
Saint-Petersburg City	0.018	-0.182 (0.066)	3.5	0.012	-0.649 (0.103)	0.7	7.0
Rep. of Dagestan	0.014	-0.178 (0.062)	3.5	0.222	-0.332 (0.093)	1.7	5.8
Rep. of Khakasia	0.013	-0.157 (0.058)	4.1	0.012	-0.264 (0.097)	2.3	3.9
Vologda Obl.	0.016	-0.157 (0.061)	4.0	0.006	-1.053 (0.372)	0.2	6.5
Astrakhan Obl.	0.032	-0.150 (0.057)	4.3	0.000	-0.193 (0.050)	3.2	2.6
Krasnoyarsk Krai	0.013	-0.145 (0.055)	4.4	0.056	-0.158 (0.064)	4.0	1.1
Amur Obl.	0.015	-0.129 (0.048)	5.0	0.000	-0.482 (0.607)	1.1	22.8
Tver Obl.	0.020	-0.128 (0.050)	5.0	0.000	-0.273 (0.165)	2.2	9.6
Kalmyk Rep.	0.056	-0.118 (0.051)	5.5	0.563	-0.637 (0.523)	0.7	18.5
Orenburg Obl.	0.052	-0.116 (0.052)	5.6	0.017	-0.342 (0.154)	1.7	15.8
Volgograd Obl.	0.022	-0.116 (0.047)	5.6	0.001	-0.185 (0.055)	3.4	3.7
Irkutsk Obl.	0.042	-0.113 (0.050)	5.8	0.001	-0.808 (0.233)	0.4	18.8
Sverdlovsk Obl.	0.064	-0.108 (0.052)	6.0	0.000	-0.238 (0.100)	2.5	3.7
Arkhangelsk Obl.	0.068	-0.104 (0.046)	6.3	0.000	-0.729 (1.185)	0.5	20.9
Kaliningrad Obl.	0.032	-0.103 (0.046)	6.4	0.007	-0.271 (0.146)	2.2	15.7
Krasnodar Krai	0.082	-0.103 (0.047)	6.4	0.000	-0.295 (0.068)	2.0	9.5
Ivanovo Obl.	0.086	-0.092 (0.045)	7.2	0.084	-0.126 (0.063)	5.1	2.6
Kurgan Obl.	0.032	-0.089 (0.039)	7.4	0.000	-0.397 (0.095)	1.4	11.5
Tyumen Obl.	0.066	-0.089 (0.046)	7.4	0.171	-0.110 (0.129)	5.9	8.8
Rep. of Northern Ossetia	0.074	-0.082 (0.040)	8.1	0.203	-0.115 (0.048)	5.6	3.8
Karachay-Circassian Rep.	0.146	-0.080 (0.044)	8.3	0.001	-0.155 (0.052)	4.1	6.0
Yaroslavl Obl.	0.104	-0.078 (0.043)	8.6	0.032	-0.247 (0.072)	2.4	6.2
Kabardian-Balkar Rep.	0.081	-0.076 (0.043)	8.8	0.002	-0.076 (0.082)	8.8	6.8
Kaluga Obl.	0.040	-0.076 (0.038)	8.7	0.000	-0.124 (0.117)	5.2	9.4
Rep. of Komi	0.119	-0.075 (0.041)	8.8	0.001	-0.165 (0.082)	3.8	6.9
Bashkortostan Rep.	0.104	-0.074 (0.042)	9.0	0.003	-0.197 (0.073)	3.2	8.9

Region	Autoregression model (1)			Threshold model (5)			
	$p(\lambda=0)$	λ	θ , months	$p(AR)$	λ^*	θ^* , months	c , %
Rep. of Karelia	0.079	-0.069 (0.036)	9.7	0.001	-0.227 (0.080)	2.7	7.9
Kursk Obl.	0.057	-0.066 (0.033)	10.1	0.066	-0.118 (0.043)	5.5	7.9
Omsk Obl.	0.104	-0.065 (0.035)	10.3	0.012	-0.197 (0.141)	3.2	13.1
Saratov Obl.	0.101	-0.064 (0.036)	10.4	0.039	-0.244 (0.106)	2.5	12.8
Smolensk Obl.	0.103	-0.064 (0.035)	10.5	0.000	-0.509 (0.418)	1.0	25.7
Vladimir Obl.	0.204	-0.063 (0.039)	10.7	0.000	-0.094 (0.051)	7.0	5.4
Stavropol Krai	0.179	-0.060 (0.036)	11.2	0.033	-0.182 (0.065)	3.4	8.5
Adygei Rep.	0.174	-0.058 (0.034)	11.6	0.318	-0.478 (0.140)	1.1	20.6
Pskov Obl.	0.200	-0.055 (0.037)	12.2	0.000	-0.146 (0.054)	4.4	5.3
Belgorod Obl.	0.216	-0.054 (0.035)	12.4	0.063	-0.108 (0.049)	6.1	10.0
Udmurt Rep.	0.196	-0.053 (0.035)	12.9	0.387	-0.179 (0.074)	3.5	8.7
Rostov Obl.	0.185	-0.051 (0.032)	13.3	0.001	-0.324 (0.077)	1.8	13.3
Nizhni Novgorod Obl.	0.160	-0.050 (0.034)	13.6	0.019	-0.183 (0.071)	3.4	6.9
Altai Krai	0.253	-0.042 (0.033)	16.2	0.000	-1.443 (0.258)	0.9	20.6
Voronezh Obl.	0.037	-0.042 (0.022)	16.3	0.170	-0.295 (0.169)	2.0	37.0
Tatarstan Rep.	0.107	-0.041 (0.025)	16.6	0.016	-0.274 (0.072)	2.2	24.0
Ryazan Obl.	0.206	-0.036 (0.030)	19.0	0.012	-0.349 (0.205)	1.6	13.4
Ulyanovsk Obl.	0.016	-0.035 (0.017)	19.4	0.002	-0.301 (0.129)	1.9	49.0
Primorsky Krai	0.310	-0.034 (0.027)	20.2	0.028	-0.091 (0.046)	7.2	23.4
Oryol Obl.	0.106	-0.033 (0.021)	20.5	0.000	-0.341 (0.184)	1.7	30.4
Bryansk Obl.	0.115	-0.032 (0.021)	21.1	0.000	-0.337 (0.193)	1.7	21.6
Kostroma Obl.	0.314	-0.032 (0.032)	21.2	0.150	-0.138 (0.063)	4.7	7.5
Sakhalin Obl.	0.287	-0.032 (0.025)	21.6	0.178	-0.099 (0.077)	6.6	41.4
Tuva Rep.	0.275	-0.032 (0.027)	21.5	0.055	-0.112 (0.052)	5.8	10.9
Jewish Autonom. Obl.	0.080	-0.031 (0.019)	21.8	0.006	-0.323 (0.110)	1.8	19.1
Kirov Obl.	0.303	-0.027 (0.024)	25.3	0.069	-0.738 (0.425)	0.5	20.4
Sakha (Yakut) Rep.	0.405	-0.027 (0.025)	25.0	0.293	-1.739 (0.757)	2.3	53.7
Penza Obl.	0.193	-0.026 (0.020)	26.1	0.023	-0.186 (0.123)	3.4	20.9
Lipetsk Obl.	0.237	-0.026 (0.021)	26.2	0.003	-0.621 (0.291)	0.7	27.3
Tula Obl.	0.286	-0.026 (0.024)	26.4	0.001	-0.649 (0.284)	0.7	21.2
Chita Obl.	0.303	-0.026 (0.023)	26.7	0.554	-0.527 (0.189)	0.9	25.9
Tambov Obl.	0.299	-0.025 (0.024)	27.1	0.000	-0.099 (0.055)	6.6	12.8
Rep. of Mordovia	0.294	-0.022 (0.020)	31.4	0.154	-0.100 (0.047)	6.6	15.7
Khabarovsk Krai	0.247	-0.021 (0.018)	33.5	0.001	-0.585 (0.199)	0.8	27.2
Murmansk Obl.	0.222	-0.018 (0.015)	38.3	0.043	-0.124 (0.062)	5.2	22.3
Mariy El Rep.	0.332	-0.014 (0.015)	48.5	0.245	-0.212 (0.116)	2.9	28.8
Chuvash Rep.	0.425	-0.010 (0.014)	65.8	0.200	-0.119 (0.055)	5.5	20.1
Magadan Obl.	0.258	-0.009 (0.008)	77.4	0.051	-0.092 (0.127)	7.2	53.7
Kamchatka Obl.	0.531	-0.006 (0.011)	114.4	0.202	-0.450 (0.155)	1.2	50.0
Moscow City	0.661	-0.003 (0.016)	228.4	0.369	-0.276 (0.135)	2.1	20.7
Mean		-0.101	6.5		-0.408	1.3	14.8
Over r with $p(AR) \leq 0.1$					-0.414	1.3	13.3

Of all the 75 regions, the hypothesis of region's integration (from the viewpoint of the strict law of one price) with the national market is rejected for 39. Thus, less than half, 48%, of regions can be deemed as integrated with the rest of the country. Taking into account the low power of the unit root test, it could be believed that there are some more integrated regions, but the test failed to support this. Adding regions with estimates that do not come a long way from the verge of significance, namely, those having p -values of λ_{r0} not more than 0.12 (9 regions), the extent of

integration would increase to 60%. This figure is not encouraging either.

Considering the geographical pattern of integration, it is seen that the major part of non-integrated regions concentrates in Central Russia. In the Central Territory (see footnote 2), there are only 4 integrated regions of 12; all regions of the Volga-Vyatka Territory are non-integrated; and there are a number of non-integrated regions that are adjacent to this “non-integration zone.” The pattern is rather surprising, as these are overwhelmingly small regions with relatively short distances between them; besides that, this part of the country has developed transport infrastructure. It can be surmised that it is just the atomistic administrative-territorial division of Central Russia that causes market fragmentation: the more regional borders and governors, the more possibilities to impede inter-regional trade and to diversify price policy across space.

Curiously, the Ulyanovsk Oblast which maintained price regulations and subsidizing as long as up to the beginning of 2001 turns out to be integrated with the national market. The time series of Moscow prices has an “almost confident” unit root with its $\lambda = -0.003$.

On the other hand, non-integrated regions are few in number in Urals and Siberia. This corroborates a finding by Gluschenko (2003) that the Asian part of Russia excluding difficult-to-access regions is more integrated than European Russia. Another evidence is the fact that all the 5 difficult-to-access regions are not integrated with the national market, a result that could be expected. It also supports findings by Gluschenko (2003, 2004a) that these regions markedly contribute to the overall disconnectedness of regional markets.

There is some similarity between the shape of the “non-integration belt” and that of the “Red Belt” reported by Berkowitz and DeJong (1999), especially in the European part of Russia. However, discrepancies are numerous, too (e.g., the difficult-to-access regions belong to the “non-integration belt”, but do not belong to the “Red Belt”). It is inconceivable that the “Red Belt” plays some role in market segmentation in the pattern obtained; this is examined in Section 4.3.

Averaging of λ across all regions yields $\bar{\lambda} = -0.101$, which corresponds to average half-life $\bar{\theta} = \theta(\bar{\lambda}) = 6.5$ months. Thus, the average speed of short-run convergence to the law of one price is rather moderate. The range of λ_{r0} is from -0.698 in the Chelyabinsk Oblast to -0.003 in Moscow; relevant θ 's vary from 0.6 to 228.4 months. The verge of statistical significance of estimates lies between θ equaling about 8 to 10 months (in terms of λ , between circa -0.080 and -0.065). However, there are three significant estimates with much longer half-lives: 16, 19, and

22 months (the mentioned Ulyanovsk Oblast just falls among them). Thus, the half-life times vary from less than one month to almost two years in integrated regions. Figure 2 plots the distribution of estimated $\{\lambda_{r0}\}$. The dashed line approximately marks the border of significance, that is, the most of estimates to its right are insignificant by the Phillips-Perron test. This line is near the median equaling -0.065 (it comes as no surprise, since the unit root is rejected for just about a half of the time series).

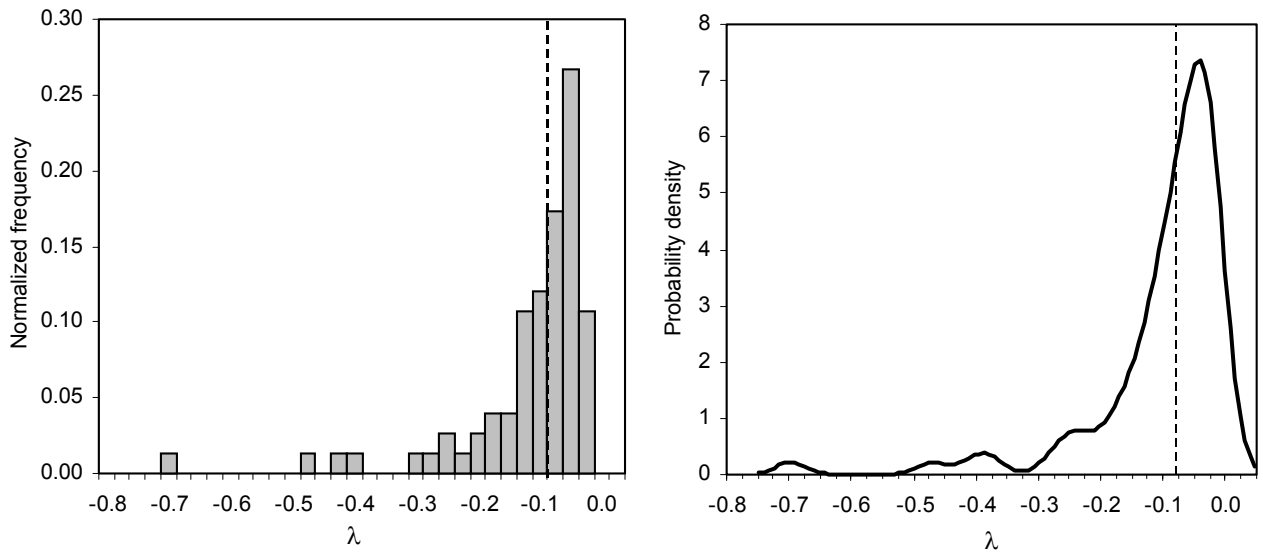


Figure 2. Distribution of estimated λ

The threshold effect is rejected for 18 regions (24% of the total), 6 of which are integrated with the national market, and 12 are not. The former case suggests that there are no impediments to trade with these regions; the latter case implies that the law of one price does not hold for the relevant regions in either (strict or weak) version.

Taking into account arbitrage inaction band, the speed of short-run convergence to the (weak) law of one price rises dramatically. Its average value increases by four times (the upper row for means in Table 1 reports values averaged over all estimates, and the lower row contains those averaged over estimates for which the TAR specification is not rejected). Figure 3 compares distributions of $\{\lambda_{r0}\}$ and $\{\lambda^*_{r0}\}$.

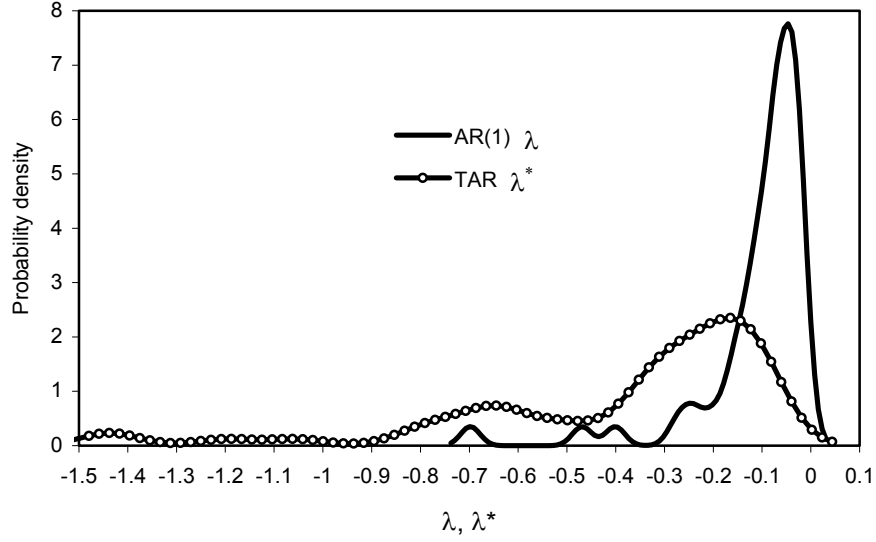


Figure 3. Comparison of distributions of estimated λ and λ^*

Both distributions in Figure 3 involve the same set of regions, namely, those with non-rejected TAR. Mean λ over such a set slightly differs from that over the entire set of 75 regions: $\bar{\lambda} = -0.108$, and the median equals -0.074 . At the same time, $\bar{\lambda}^* = -0.414$ and the median of λ^* is equal to -0.273 . The relevant average half-lives are 6.1 and 1.3 months respectively. The TAR model yields θ^* less than or equal to 1 month for 19 regions, and the maximum magnitude of the half-life time equaling 8.8 months.

Along with this, the thresholds representing arbitrage transaction costs are high. In the Magadan Oblast, the threshold is about 54%. This is not surprising for difficult access to the region. The surprising thing is an enormous threshold in the Ulyanovsk Oblast, where its value is 49%. It may be concluded that this is just a result of prolonged (namely, covering the entire period under consideration) maintaining price regulations, subsidizing and other protectionist measures by the government of this region. In 14 regions, the threshold varies from 20 to 30%, while there are only 7 regions with thresholds below 5%. For reference, dealing with the food component of the national CPI, Obstfeld and Taylor (1997) find the threshold to vary in the range of 0.6% to 4.3% across US cities (though, this result is not too representative, since the case at hand is only 4 cities). The entire distribution of estimated $\{c_{r0}\}$ is presented in Figure 4.

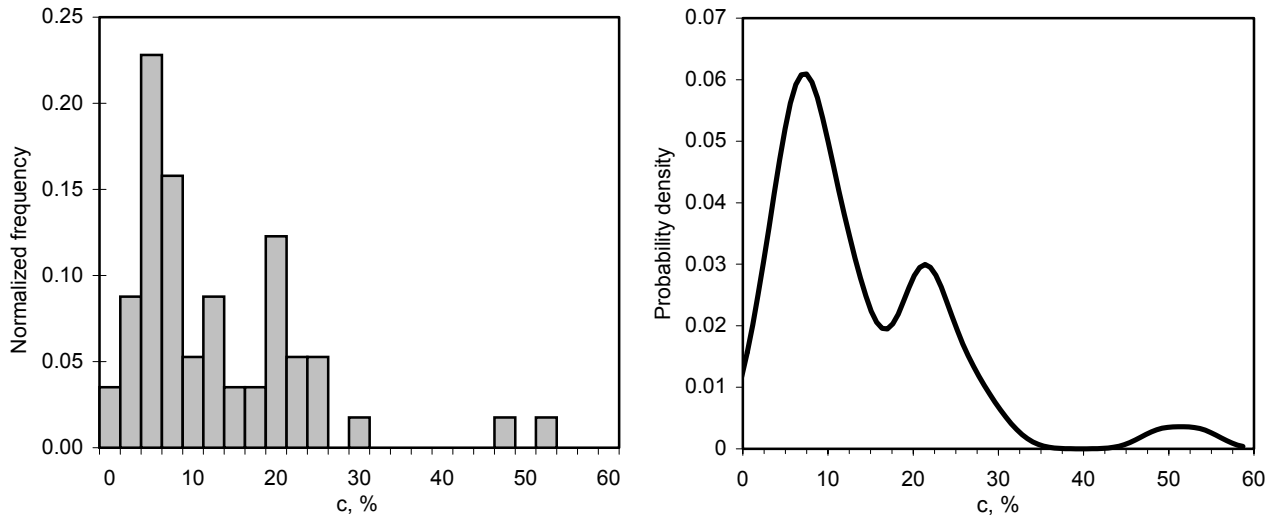


Figure 4. Distribution of estimated thresholds c

The mean threshold equals 13.3% of the average Russian price (this mean is calculated by Formula (6) as $c(\overline{C}_{(+)})$; directly averaging yields a higher value, $\bar{c} = 13.7\%$); and the median equals 10%. Thus, in a half of 57 regions where the threshold effect takes place, arbitrage transaction costs exceed 10% of the average national price.

4.2. Analyzing long-run price convergence

Having considered the state of market integration in 1994-2000, the evolution of integration during this period is turned to. As less than a half of Russian regions can be deemed as integrated, the question rises whether the rest regions tend to be more integrated, that is, whether there is long-run price convergence among them. To answer this question, price distribution dynamics is analyzed in this section.

The first issue is that of σ -convergence. Figure 5 plots the dynamics of price dispersion – measured as σ_t , the standard deviation of (log) prices normalized to the Russian average, $\{P_{r0}\}$. The trajectories of σ_t are estimated over all the 75 regions under consideration, and over integrated and non-integrated regions separately. For the latter, the trajectory is also estimated with the elimination of the 5 regions with difficult access in order to shed light on their impact. For comparability, σ_t for each region group is computed with the use of the mean over all regions rather than that over a given group. That is, the price dispersion is measured relative the

whole of Russia; hence, it is not a within-group dispersion. With this, the price dispersion over all Russian regions is a weighted average of dispersions over integrated and non-integrated regions: $\sigma_{(all)t} = 0.48\sigma_{(int)t} + 0.52\sigma_{(non-int)t}$, the weight being the share of the group in the total number of regions.

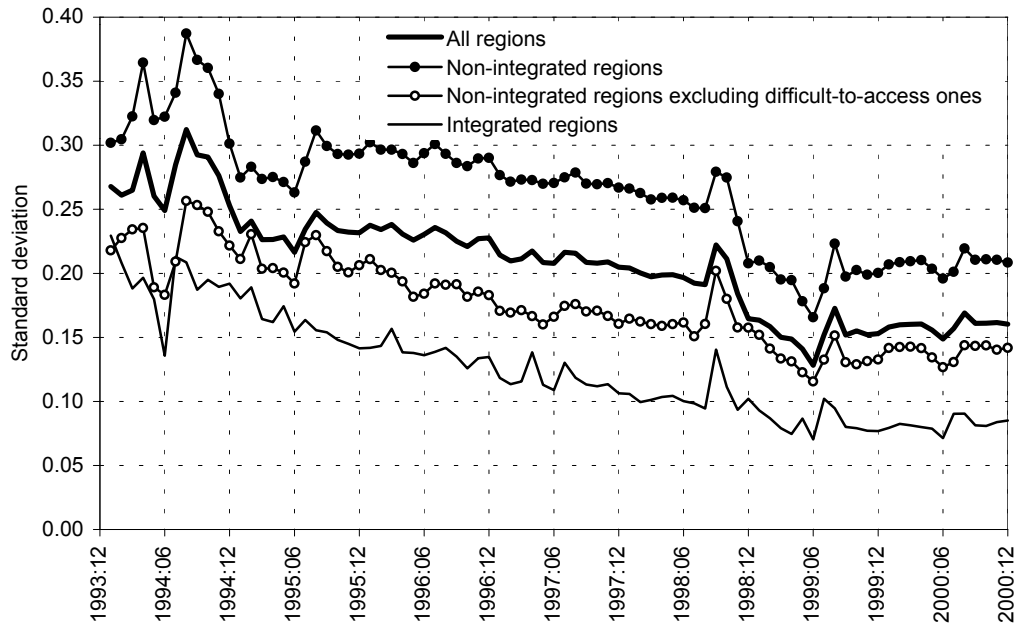


Figure 5. The evolution of price dispersion

The figure provides clear evidence of σ -convergence among non-integrated regions. Comparing relatively distant points in time (so that to eliminate random disturbances), Formula (3) holds for trajectory for all non-integrated regions and that for those excluding difficult-to-access regions. Being treated as time series, both the trajectories prove to be trend stationary with a highly significant negative trend factor, thus also suggesting σ -convergence.

Comparing trajectories for non-integrated regions with and without difficult-to-access ones, it is seen that the latter sufficiently contribute to the overall price dispersion. However, the price gap between difficult-to-access regions and the rest of Russia also tends to decrease with time.

Figure 5 demonstrates a caveat which is not noticeable when each time series is separately dealt with. The point is that the trajectory for integrated regions looks strange. Since each of individual time series for this set of regions is recognized as stationary, its variance has to be near-constant. Hence, $\sigma_{(int)t}$ is expected to be near-constant as well. But the aggregated trajectory

conflicts with such an expectation, suggesting decreasing variance. From this it follows that some of time series, albeit satisfying the law of one price test (1), are not in fact stationary, but have a trend towards price parity. To recognize this, a subtler econometric tool than simple AR(1) model is necessary.

One more caveat is a sharp shock in the middle of the second half of 1998, which is due to the August 1998 financial crisis in Russia. Figure 5 provides a visual impression that this shock could cause structural breaks in the price time series. Thus, for more thorough analysis, (1) should be augmented for such a break.

Let us turn to the behavior of the entire cross-section distribution of prices. At first, the issue of interest is changes in its shape over time. The density functions are estimated with the use of Formula (7). Three region samples are dealt with: all 75 regions, $R_1=R=75$; non-integrated regions, $R_2=39$; and integrated regions, $R_3=36$. Bandwidth h varies across these samples, but it is put to be uniform across t for each given sample; the value of h is taken to be about Silverman's (1986) bandwidth averaged over involved points in time. These points are 1994:01 and each December of 1994 through 2000. The estimated densities are reported in Figure 6 for all and non-integrated regions and for selected time points. Appendix B reports plots for each involved point in time as well as for integrated regions.

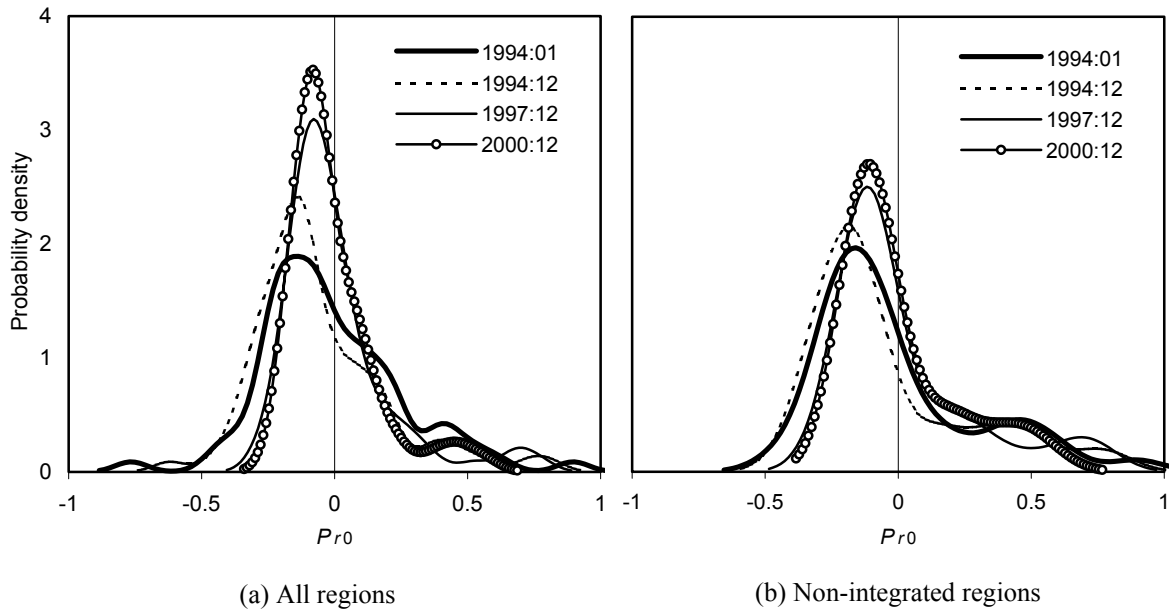


Figure 6. Selected estimates of cross-region price distributions

The densities in the figure reveal several features in the distribution shape dynamics over the 7-year period, January 1994 to December 2000. Turning to the sample of all regions, Figure 6(a), the mode of the distribution is shifting from negative P towards zero, or, in terms of prices, from the cost of the 25-food basket below the Russian average towards the national average. Along with this, the left-hand tail of the distribution shortens with time. At the same time, the long and heavy right-hand tail is persistent throughout the period, and prevents the distribution from becoming symmetric by the end of the period. However, the right tail did somehow shorten. Referring to Figure B1 in Appendix B, it is seen that the most prominent shift occurs between 1997:12 and 1998:12. Not reported monthly estimates evidence that this shift happened shortly after the August 1998 financial crisis. For the most part, prices responsible for the right-hand tail are those in the difficult-to-access and Far-Eastern regions. The share of foods imported from abroad was smaller in this part of the country (as well as in the Asian part a Russia in general) than in its Western part. Therefore the sharp devaluation of ruble in August 1998 lowered (relative) prices in most of the “expensive” regions. In this way, the 1998 crisis narrowed the price gap between the Far East and the rest of Russia.

Being a subsample of the above-considered sample, non-integrated regions demonstrate some similar properties of distributions for them in Figure 6(b), namely, the shift of the mode towards zero over time and persistent heavy right-hand tail. Two main differences are as follows. Firstly, there is no long left-hand tail in the distributions for non-integrated regions. It is little wonder, since this tail is due to “cheap” integrated regions. Secondly, the right tail of the distributions is heavier. Again, it comes as no surprise, as it is non-integrated “expensive” regions that concentrate in this tail. The distribution for non-integrated regions has the following statistics in 1999:12 and 2000:12 as compared to those for Russia as a whole (in parentheses): the mean: 0.009 and 0.000 (−0.009 and −0.010), the median: −0.076 and −0.088 (−0.037 and −0.060), the standard deviation: 0.200 and 0.209 (0.153 and 0.160).

Figure 6 gives a visual impression of small secondary modes in the right and left tails of distributions for some years, so seemingly suggesting the presence of price convergence clubs. But this is not the fact. This “modes” occur when prices in two (sometimes, three) regions accidentally become close. Since the observations are sparse in those parts of the distribution, the estimator given by Formula (7) smoothes such an “outlier” in a histogram into a small mode. Moreover, testing for multimodality in the manner of Silverman (1986), three- and bimodality of the distributions with seeming secondary mode(s) is confidently rejected.

At last, a few words about the distribution for integrated regions (Figure B3 in Appendix B). It tends towards a symmetric one; its mean and median are close to one another and to zero (they are respectively -0.029 and -0.030 in 1999:12, and -0.020 and -0.042 in 2000:12). The distribution is much narrower than that for Russia as a whole: the standard deviation of the former equals 0.083 , while that of the latter is equal to 0.160 . Besides that, the distribution for integrated regions tends to normality. For example, the hypothesis of normality has significance of 71% (by the Jarque-Bera statistic) in 2000:12.

Overall, turning back to non-integrated regions, it can be concluded that the distribution of their prices has a tendency of coming closer to the distribution of integrated regions. However, this does not hold for the right-hand tail which remains long and heavy. Since the end of 1996, almost all of regions in the tail are remote ones except for Moscow. In 1999 and 2000, the rightmost part of the tail, where prices are more than 15% above the national average, includes all the 5 difficult-to-access regions, the Khabarovsk Krai and the Primorsky Krai (the almost easternmost regions of Russia), and Moscow. Their prices – at least, all of them – can be hardly believed to lower so that the right-hand tail will become similar to the left-hand one. Thus, taking account of this geographical feature, the changes in the distribution shape evidence long-run price convergence among non-integrated regions. The distribution is unimodal, so suggesting the absence of price convergence clubs.

Having obtained the sequence of price distributions, the next issue is that of a law of motion. The evolution of the entire cross-region distribution of prices is modeled by the transition probability matrix; see Formula (8). It can be also interpreted as a characterization of (absolute) mobility of regions in terms of prices.

In this analysis, absolute values of prices normalized to the Russian average, p_{rt}/p_{0t} , are used rather than their logarithms. The range of possible values of regional prices is divided into 5 price classes: prices more than 15% below p_{0t} , the national average, $[0, 0.85)$; prices 15% to 5% lower than p_{0t} , $[0.85, 0.95)$; prices between 5% lower and higher than p_{0t} , $[0.95, 1.05)$; prices 5% to 15% higher than p_{0t} , $[1.05, 1.15)$; and prices 15% and more above p_{0t} , $[1.15, \infty)$.¹² The reasoning behind such a discretization is as follows. First, the price classes are chosen so that they have the same size (certainly, except for the extreme ones). Second, the grid is constructed symmetrically about 1, the national average. Third, the discretization is fitted so that no one class

¹² In logarithms, such a grid is slightly asymmetric, being equivalent to $(-\infty, -0.16)$, $[-0.16, -0.05)$, $[-0.05, 0.05)$, $[0.05, 0.14)$, and $[0.14, \infty)$.

degenerates into an empty one during the covered period. And forth, the borders of the classes are wanted to be expressed in round numbers. After experimenting with various discretizations, the reported grid was chosen as the best one. The time points are the same as for the kernel density estimations above.

Table 2 reports the number of regions in each price class in a given point in time as well as the total number of regions falling into each class during the entire period. In other words, a column of this table is a transposed (non-normalized) probability vector \mathbf{f}_t (see Section 2.3).

Table 2. Distribution of regions by price class

Price class	Number of regions in a given price class								Total
	1994:01	1994:12	1995:12	1996:12	1997:12	1998:12	1999:12	2000:12	
< 0.85	24	27	18	13	9	8	6	3	108
0.85-0.95	16	20	21	30	28	24	27	37	203
0.95-1.05	9	8	14	11	19	24	26	15	126
1.05-1.15	9	6	5	7	7	8	8	12	62
> 1.15	17	14	17	14	12	11	8	8	101

From considerations of robustness of results, the transition probability matrix is estimated in two ways. The first uses information only on price transition of regions between the initial and final points of the time span concerned, i.e., the matrix describes one 7-year transition from 1994:01 to 2000:12, $\mathbf{M}(\tau) = \mathbf{M}(7 \text{ years})$. The second way makes use of information on transitions within 1994-2000; however, the more distant is a transition in time, the lesser importance is attached to it. That is, the estimator of the transition probability matrix is a weighted average of year-to-year estimates: $\mathbf{M}(\tau) = \overline{\mathbf{M}}(1 \text{ year}) = w(\mathbf{M}(1994:01 \text{ to } 1994:12)/7 + \mathbf{M}(1994:12 \text{ to } 1995:12)/6 + \mathbf{M}(1995:12 \text{ to } 1996:12)/5 + \mathbf{M}(1996:12 \text{ to } 1997:12)/4 + \mathbf{M}(1997:12 \text{ to } 1998:12)/3 + \mathbf{M}(1998:12 \text{ to } 1999:12)/2 + \mathbf{M}(1999:12 \text{ to } 2000:12)/1)$, where $w (\approx 0.386)$ is a normalizing factor making the weights to sum to unity.

As described in Section 2.3, transition probability matrices can be used for estimating a long-run limit of the price distribution, the ergodic distribution; see Formulae (9) and (10). The estimated transition probability matrices along with the resulting ergodic distributions are presented in Table 3. Its upper panel contains $\mathbf{M}(\tau)$ estimated from the one-step transition, from 1994:01 to 2000:12. The bottom panel contains the transition probability matrix estimated by averaging the one-year transitions across years; Appendix C reports all the 7 yearly matrices

used for the averaging.

Table 3. Transition probability matrices and ergodic distributions

Initial price class	Probability of transition to the price class of				
	< 0.85	0.85-0.95	0.95-1.05	1.05-1.15	> 1.15
Transition from 1994:01 to 2000:12, $\mathbf{M}(7)$					
< 0.85	0.125	0.708	0.167		
0.85-0.95		0.875	0.125		
0.95-1.05		0.333	0.333	0.222	0.111
1.05-1.15		0.333	0.333	0.333	
> 1.15		0.059	0.118	0.412	0.412
Ergodic distribution		0.707	0.179	0.080	0.034
Weighted average of 7 one-year transitions, $\overline{\mathbf{M}}(1)$					
< 0.85	0.497	0.479	0.024		
0.85-0.95	0.078	0.728	0.182	0.010	0.003
0.95-1.05	0.026	0.301	0.494	0.151	0.028
1.05-1.15		0.034	0.353	0.586	0.028
> 1.15			0.025	0.094	0.881
Ergodic distribution	0.081	0.438	0.254	0.126	0.101

Note: Cells showing 0 are left blank.

In matrix $\mathbf{M}(7)$, price mobility is rather high. Among the diagonal elements, which represent the probability of staying in the same state, there is the only one (for moderately low – 0.85 to 0.95 – prices) with high value. The most probability is to transit to the price range of 0.85 to 1.05 of the national average, except for regions from the highest price class (above 1.15). There are 9 forbidden transitions, i.e., those having the zero probability, all transitions to very low prices (below 0.85) falling among them. The pattern of intra-distribution mobility provided by matrix $\overline{\mathbf{M}}(1)$ is more conservative. All the diagonal elements equal to roughly 0.5 and more; the probability of persisting the highest prices is about 0.9. At the same time, the number of the forbidden transitions falls to 5. This inconsistency between transition matrices estimated over single and multiple periods is a well-known problem discussed, e.g., by Singer and Spilerman (1976), and that is why both methods of estimating \mathbf{M} are applied.

Nevertheless, both matrices have the same qualitative feature. They suggest that regions tend, in general, to transit to the price range of 0.85 to 1.05 of the national average. That is, regions with lower prices tend to shift to higher ones, and regions with higher prices tend to shift to lower ones. This feature evidences in favor of long-run price convergence.

The bottom lines in each panel of Table 3 contain ergodic distributions; 150 to 180 iterations (exponentiations) have been sufficient for matrices to converge to them. The distributions are portrayed in Figure 7 along with the actual distribution in the end of the period under consideration (in fact, it the last but one column of Table 2 normalized to $R = 75$).

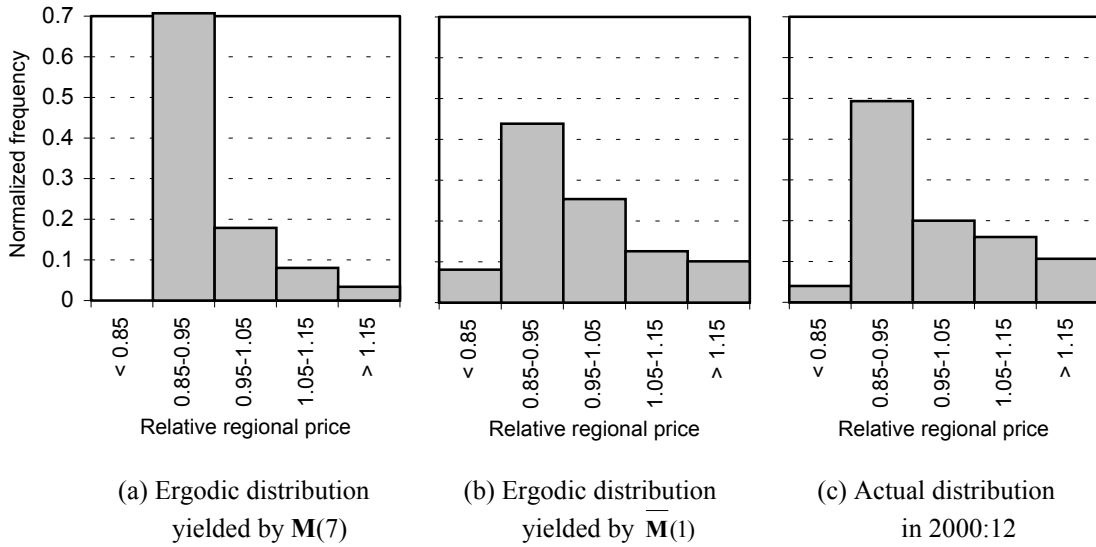


Figure 7. Ergodic price distributions

Both distributions are non-symmetric with a heavy right-hand tail, thus suggesting the long-run persistence of regions with very high prices. Furthermore, both distributions are unimodal, thus suggesting the absence of price convergence clubs in the long run.

Along with this, the distributions differ from one another, as do the matrices generating them. That obtained from 7-year transition matrix $\mathbf{M}(7)$ is much thinner, the low-price regions being a vanishing class. The distribution yielded by one-year transition matrix $\mathbf{M}(1)$ is rather close to the actual price distribution in 2000:12. However, it is not inconceivable that these differences in many respects are due to the discretization used to construct the transition probability matrices that is, in fact, arbitrary (and, needless to say, very crude).

4.3. Impediments to market integration

This section aims to reveal reasons behind the pattern of integration considered in Section 4.1. Thus, the variables to be explained are estimates from Table 1. The following additional variables representing potential impediments to integration are used (for more detailed

description of these, see Gluschenko, 2004a):

Distance, the average (log) distance from the region to the rest regions, normalized to the average inter-regional distance over all regions; the distance matrix used is that compiled by Alexei Abramov, Novosibirsk State University;

Difficult access to the region, a dummy equaling 1 if so, and 0 otherwise;

Infrastructure, measuring the quality of a region's transport infrastructure from 0 to 1 so that the less the index, the better infrastructure; the index is drawn from Matiyasevich *et al.* (1998);

Regional transport tariff, the average over 1994-2000 of the (log) annual indices of tariffs for freight shipments in the region; the raw data come from Goskomstat (2000, 2002);

Distribution costs, the average over 1994-1997 (the data for 1998-2000 are not available) of the (log) annual ratio of trading costs to retail sales in the region; the raw data come from Goskomstat (1996b, 1998);

Price regulations, the (log) proportion of goods and services with regulated prices in the region in the first quarter of 1996 according to Goskomstat's survey;

Subsidy, the (log) production subsidies as a proportion of the regional budget expenditures in 1995 according to Goskomstat's survey;

Red Belt, a dummy equaling 1 if the region belongs to the "Red Belt", and 0 otherwise; the "Red Belt" is defined according to Berkowitz and DeJong (1999) as regions that voted in favor of the Communist Party of the Russian Federation during the 1996 presidential elections;

Organized crime, the proportion of the regional economy controlled by criminal groups in 1995-1996; the quantification from 0 to 1 is drawn from Brown and Earle (2000);

Crime rate, the average over 1994-2000 of the (log) annual number of registered crimes in the region per 100,000 of the population, normalized to the index for entire Russia; the raw data come from Goskomstat (2001).

Binary choice model (11) – with the left-hand side variable equaling 1 if the law of one price is rejected for a given region, and 0 otherwise – suggests that the only variable accounts for non-integration, the subsidy; all the rest variables are insignificant. However, this evidence is very uncertain, since it is rejected with robustly estimated standard errors in the logit model; the probit model does not support the finding at all. The estimates are as follows:

Model	Coefficient on Subsidy	Standard error / p -value	Robust standard error / p -value	Pseudo R^2
Logit	1.900	1.071 / 0.076	1.434 / 0.185	7.6%
Probit	0.816	0.516 / 0.114	0.668 / 0.222	6.3%

Neither the ordinary linear model (11') – the left-hand side variable in which is the Phillips-Perron test p -value – supports the significance of the subsidy variable. In this model, only the difficult access and transport infrastructure are significant, the latter being rejected with robust standard errors. Nevertheless, the R^2 statistic is not too low, equaling 22.3%. The estimates are as follows:

Variable	Coefficient	Standard error / p -value	Robust standard error / p -value
Difficult access	0.254	0.058 / 0.000	0.065 / 0.000
Infrastructure	0.239	0.098 / 0.017	0.201 / 0.239

Thus, the probability of non-integration rises if the region is a difficult-to-access one. This adds nothing to our knowledge, as it is known that all difficult-to-access regions have been found as non-integrated. Another estimate is more informative, suggesting that the worse transport infrastructure in the region, the less probability of region's integration with the national market. Contrary to the expectations the Red Belt turns out to be insignificant in both above regressions. This implies that belonging of the region to the "Red Belt" accounts for neither its non-integration with the national market nor the probability of being non-integrated.

Table 4 summarizes results of analyzing thresholds \hat{c}_r , i.e., estimated arbitrage transaction costs, with the use of model (12). Hereafter, robust standard errors are reported. The regression is run over regions for which the threshold effect is not rejected ($p(\text{AR}) \leq 0.1$ in Table 1).

Table 4. Impact of specific frictions on the total effect of all impediments to inter-regional trade

Variable	Coefficient	Standard Error	p -value
Russia as a whole (57 observations)			
Difficult access	26.235	9.020	0.005
Distance	5.544	2.308	0.020
Subsidy	8.772	3.206	0.008
Red Belt	4.763	2.317	0.045
Excluding difficult-to-access regions (55 observations)			
Distance	3.857	2.418	0.117
Subsidy	8.130	3.188	0.014
Red Belt	4.636	2.281	0.047

The difficult access is to be expected to sufficiently raise barriers to trade. But other findings are not so trivial. The fact that distance contributes to arbitrage transaction costs is of no surprise by itself. But this forces to treat the pattern of market integration with more caution, implying that the rejection of integration may be too severe, as it does not allow for this “natural” impediment to trade, the spatial separation of regions. It is obvious that the distance and the difficult access are correlated, and the regression in the lower panel of Table 4 eliminates this. Its results suggest that the distance estimate is not robust to such a multicollinearity, while the rest two estimates are. Albeit the distance estimate becomes insignificant, it does not go far from the 10% significance level. And so, it should not be ruled out that distance does matter.

While merely a hint of some role of subsidizing has been obtained above, here its importance becomes pronounced. Subsidizing does contribute to disconnectedness of regional markets, widening inter-regional price gaps. The “Red Belt” appears, at last, as an impediment to market integration. Although it does not impact directly on region’s integration with the national market, it acts implicitly, through increasing inter-regional barriers.

At last, linkages between the AR and TAR estimates were analyzed. The regression of the LLR test p -value on the Phillips-Perron test p -value ($p(\text{AR})$ and $p(\lambda=0)$, respectively, from Table 1) over all 75 regions yields the point estimate of 0.333 with the standard error equaling 0.009 (significance at the 0.1% level); $R^2=12.6\%$. This implies that the higher probability of non-rejection of unit root in a time series, the higher probability that the threshold effect will be rejected as well (in other words, that the law of one price is violated in either version). The regression of estimated threshold \hat{c}_r on the non-integration/integration (0/1) variable has $R^2=10.4\%$. The estimate equals 6.693, having the standard error of 2.660 (p -value is equal to 0.015). Hence, barriers to trade are higher in non-integrated regions, as would be expected. One more regression is that of threshold \hat{c}_r on half-life time $\hat{\theta}_r$, yielding $R^2=55.0\%$. The estimate has the value of 0.633 and the standard error of 0.050 (yielding zero p -value). Thus, the lower the speed of short-run convergence to the law of one price, the higher the threshold (or, conversely, the higher the threshold, the lower the speed). This can be interpreted as follows. Assuming the data generating process to be AR(1), while it is in fact a TAR process, one understates the convergence speed, since the process actually converges to the external borders of the arbitrage inaction band rather than to zero. Hence, the result evidences that the threshold effect is truly present in the time series used.

The results reported in this section are not too rich, especially as compared to those of the cross-sectional analysis in Gluschenko (2004a). A possible reason may be that the AR and TAR estimates – used as variables to be explained – average the behavior of prices over the whole period of 1994:01-2000:12. On the other hand, a number of explanatory variables are time averages as well. Thus, the pattern becomes very rough, losing many sufficient details (in fact, the transitional nature itself of the Russian economy). And so, regressing a crude indicator on a set of other such indicators fails in revealing a number of more subtle dependencies.

5. CONCLUSIONS

Using the cost of the basket of 25 basic food goods as the price representative, the state and trends of market integration in Russia has been analyzed. The results obtained evidence rather poor market integration over 1994-2000: with the strict law of one price as the benchmark, only 50% to 60% of Russian regions can be deemed as integrated with the national market. Besides that, barriers to inter-regional trade are found to be rather high; being integrally quantified as arbitrage transaction costs, they equal on average 13% of the price for Russia as a whole. Nonetheless, encouraging evidence is found as well, which is long-run price convergence. This implies that integration of the Russian market is tending to improve.

Besides that, it is inconceivable that the obtained pattern of integration overstates shortcomings of the Russian market. In particular, the 1998 structural break could bias statistical inferences towards non-rejection of unit root in some instances. However, it is not this fact that is of first importance.

In this paper, the methodology of time series analysis is used that is usually applied to long-standing market economies for testing the law of one price and PPP. Albeit results yielded by it look reasonable, there is a serious caveat. A tacit assumption behind this methodology is that the nature of price behavior does not change much over time. But in transition economies, such an assumption is hardly valid; and what is more, changes in the nature of price behavior are one of aspects of the transition itself. As a result, the time series that are recognized as stationary with the use of the conventional AR model, (1), taken jointly, provide contradictory evidence (see Figure 5). As for the threshold model, changes in the nature of price behavior imply that the thresholds can be time-dependent. If so, TAR (5) turns out to be too crude to model price behavior.

Thus, analyzing market integration in transition economies (as well as in markets experiencing sufficient changes in the nature of price behavior, e.g., in the Eurozone after introducing the cash euro) needs methodologies that would be more adequate to their features. Such a methodology is put forward in a subsequent study presented in Gluschenko (2004b). Its application does not change the above pattern fundamentally; however, the pattern becomes more comprehensive and clear, the results obtained through different analyses proving to be consistent.¹³ It follows herefrom that the traditional tools of time series analysis can be, in principle, applied to examining the price behavior in transition economies; however, they are impotent in revealing fine details peculiar to transition processes.

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¹³ If merely the structural break were taken into account, the share of integrated regions would indeed increase to 63%. Incorporating long-run convergence into time-series modeling (with the use of nonlinear, asymptotically subsiding trend), of 36 regions recognized as integrated in this paper, 15 are classed with those tending to integration (while one is found non-integrated without such a trend); of 39 non-integrated regions, 7 are classed with integrated ones, and 18 are classed with ones tending to integration. Overall, 36% of regions are found integrated, 44% are found to tend to integration, and only 20% are recognized as non-integrated without trend to integration.

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APPENDIX A: Upper and lower arbitrage thresholds

Let z_{rs} and z_{sr} be absolute values of arbitrage transaction costs (per physical unit of a given good) in the cases of delivering the good from regional market r to market s , and from s to r , respectively, and p_r and p_s be prices of the good in regions r and s . An inflow of the good to r from s is unprofitable when $p_r \leq p_s + z_{sr}$; an outflow from r to s is unprofitable when $p_s \leq p_r + z_{rs}$. Hence, arbitrage between markets r and s does not operate if

$$p_s - z_{rs} \leq p_r \leq p_s + z_{sr}, \quad (\text{A1})$$

or, equivalently,

$$1/(1 + z_{rs}/p_r) \leq p_r/p_s \leq 1 + z_{sr}/p_s. \quad (\text{A2})$$

Commonly, basing on the Samuelson (1954) iceberg model of transportation costs, it is taken $z = \delta p$, i.e., $z_{rs}/p_r = z_{sr}/p_s = \delta_{rs}$. Thus, the arbitrage inaction band is

$$-C_{rs} \leq P_{rs} \leq C_{rs}, \quad (\text{A3})$$

where $C_{rs} = \ln(1 + \delta_{rs})$, $P_{rs} = \ln(p_r/p_s)$.

However, the assumption of proportional transportation costs (and transaction costs in general) is hardly realistic. For example, if a good were twice as expensive in s than in r , shipping the good from s to r would be twice as expensive, too, than shipping from r to s . But, according to experts, the actual share of the price-dependent component of transportation costs does not exceed a few percents of the total expense for shipping an ordinary cargo (such as foods and the like).

Hence, it is much more realistic to assume transaction costs to be constant for a given good and location pair regardless of the direction of transportation, i.e., $z_{rs} = z_{sr}$. Taking p_s as a base for calculating the percentage of transaction costs, $c_{rs} = z_{rs}/p_s$, Formula (A1) rearranges to

$$1 - c_{rs} \leq p_r/p_s \leq 1 + c_{rs}, \quad (\text{A4})$$

or, in the logarithmic form,

$$C_{(-)rs} \leq P_{rs} \leq C_{(+)rs}, \quad (\text{A5})$$

where $C_{(-)rs} = \ln(1 - c_{rs})$ and $C_{(+)rs} = \ln(1 + c_{rs})$.

It is easily seen that the upper and lower arbitrage thresholds, $C_{(-)rs}$ and $C_{(+)rs}$, are related by relationships

$$C_{(-)rs} = \ln(2 - \exp C_{(+)rs}), \quad C_{(+)rs} = \ln(2 - \exp C_{(-)rs}). \quad (\text{A6})$$

It is worth noting, however, that the assumption $z_{rs} = z_{sr}$ is a simplification as well when components of transaction costs other than transportation costs are taken into account. For example, regional authority could prohibit exporting foods from region r , while there are no such impediments in region s . In such a situation, $z_{rs} > z_{sr}$. Hence, in the general case, the barrier between r and s can be anisotropic, $c_{rs} \neq c_{sr}$; as a result, a functional relationship between $C_{(-)rs}$ and $C_{(+)rs}$ – like (A6) – is lacking.

APPENDIX B: Non-parametric estimates of cross-region price distributions

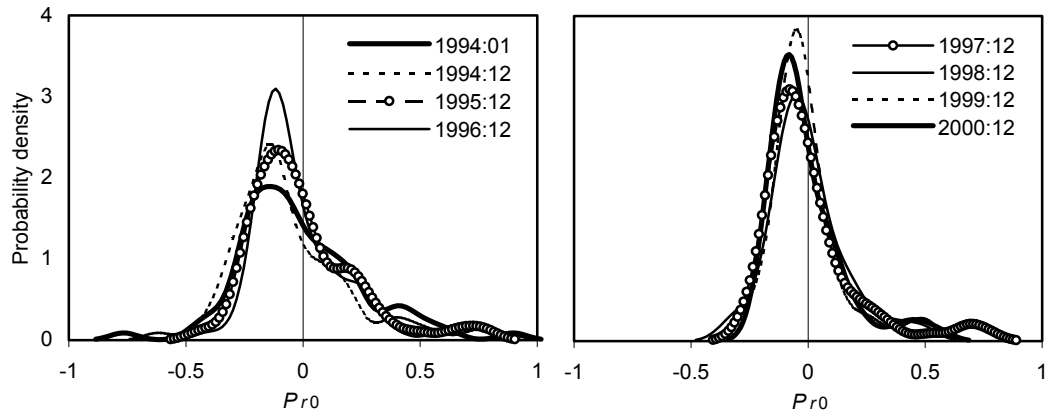


Figure B1. All regions

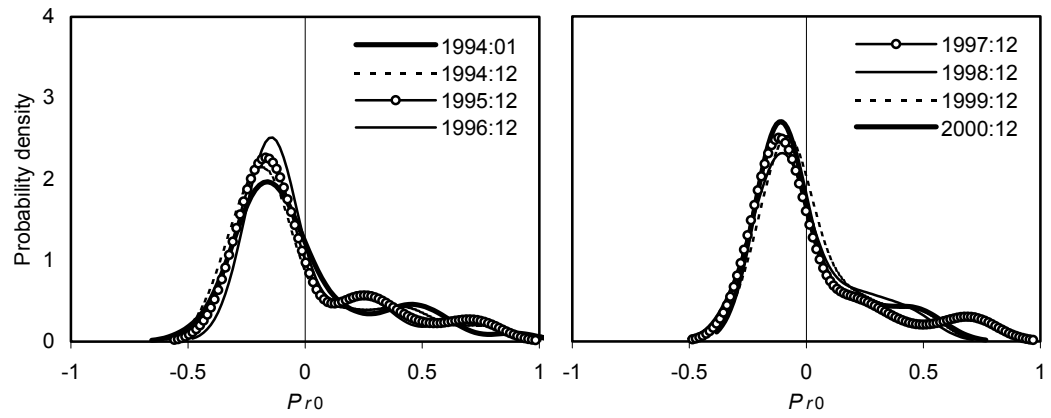


Figure B2. Non-integrated regions

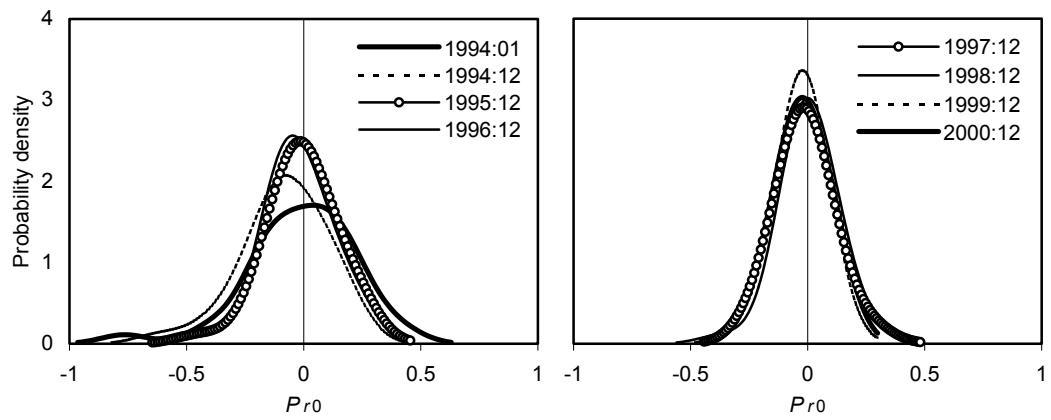


Figure B3. Integrated regions

APPENDIX C: Yearly transition probability matrices

Initial price class	Probability of transition to the price class of				
	< 0.85	0.85-0.95	0.95-1.05	1.05-1.15	> 1.15
1994:01 to 1994:12					
< 0.85	0.708	0.292			
0.85-0.95	0.500	0.438	0.063		
0.95-1.05	0.222	0.333	0.111	0.111	0.222
1.05-1.15		0.333	0.111	0.444	0.111
> 1.15			0.294	0.059	0.647
1994:12 to 1995:12					
< 0.85	0.556	0.407	0.037		
0.85-0.95	0.150	0.450	0.200	0.150	0.050
0.95-1.05		0.125	0.500	0.125	0.250
1.05-1.15			0.500	0.167	0.333
> 1.15			0.143		0.857
1995:12 to 1996:12					
< 0.85	0.556	0.444			
0.85-0.95	0.095	0.857	0.048		
0.95-1.05	0.071	0.214	0.571	0.143	
1.05-1.15		0.200	0.400	0.400	
> 1.15				0.176	0.824
1996:12 to 1997:12					
< 0.85	0.462	0.462	0.077		
0.85-0.95	0.100	0.633	0.267		
0.95-1.05		0.273	0.636	0.091	
1.05-1.15			0.429	0.571	
> 1.15				0.143	0.857
1997:12 to 1998:12					
< 0.85	0.667	0.222	0.111		
0.85-0.95	0.071	0.643	0.286		
0.95-1.05		0.211	0.579	0.211	
1.05-1.15			0.571	0.429	
> 1.15				0.083	0.917
1998:12 to 1999:12					
< 0.85	0.625	0.375			
0.85-0.95		0.708	0.292		
0.95-1.05	0.042	0.292	0.583	0.083	
1.05-1.15			0.625	0.375	
> 1.15				0.273	0.727
1999:12 to 2000:12					
< 0.85	0.333	0.667			
0.85-0.95	0.037	0.852	0.111		
0.95-1.05		0.385	0.423	0.192	
1.05-1.15			0.125	0.875	
> 1.15					1.000

Note: Cells showing 0 are left blank.