THE LAW OF ONE PRICE AND REGIONAL PRICE CONVERGENCE IN UKRAINE

by

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Abstract

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The study of 30 nominal retail prices for 26 regions in Ukraine for the period 1997:01-2002:12 shows that the law of one price holds for relative prices to the average Ukrainian level when transaction costs are taken into account in estimation procedure. Even though, panel unit root tests are unable to reject the null hypothesis of nonstationarity in relative prices, Band-TAR model estimation results suggest that regional prices do converge to the average Ukrainian level. Food goods have higher speed of convergence and lower threshold levels than nonfood commodities and services with half-lives ranging from 1 to 2 months and threshold levels from 3% to 15%. In accordance with other studies, speed of convergence for services is three time slower than convergence for food group.

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GLOSSARY

LOP - the law of one price

PPP – purchasing power parity

Half-life -the number of periods it takes to reduce by half the magnitude of a deviation from the equilibrium caused by an individual shock

CPI – Consumer Price Index

RPI – Retail Price Index

Chapter 1

INTRODUCTION

The intranational law of one price (henceforth LOP), as the basis of intranational 'purchasing power parity' (PPP) for individual goods is an empirical proposition that in competitive markets, free of transportation costs and official barriers to trade (such as tariffs), identical goods sold in different locations must sell for the same price when their prices are expressed in terms of the same currency (Krugman and Obstfeld, p. 395).

Though the law of one price remains one of the most important building blocks of international economics, empirical investigations generally failed to prove its existence. This failure to detect price convergence among different countries undermined confidence in a wide range of open-economy macro models that assumed some version of the law of one price. "The difficulties researchers had in rejecting a random walk model for PPP deviations on modern floating rate data was something of an embarrassment. Every reasonable theoretical model suggests that there should be at least some temporary component to PPP deviations."(Rogoff, 1996, p.655) Therefore, there is great interest in determining factors that cause persistent deviations from the LOP.

In the face of the negative conclusions of empirical literature there has been an increasing incentive for researchers to use new econometric techniques and data to investigate the topic of price convergence. One of the tools aimed to reduce the possibility of rejection of LOP hypothesis is to use an intranational rather than an international context.

Typically, there is much more trade in goods between regions and cities within a country, than between countries – many barriers and distortions to trade are usually much smaller inside a country[1]. Thus, regional data allows the elimination of two major sources of deviation from the LOP: exchange rates and trade barriers. Moreover, intranational violations of the LOP may matter more than international ones because the former are more likely to involve substantial regional quantity imbalances and, thus, resource misallocations, which is of particular concern to national policy makers. In addition, differences in cost-of-living between regions are important because the systems of transfer payments may not take them into account.

In case of transitional economies, the issue of the law of one price is of additional importance, a fiture associated with the fact that nearly all transition countries have passed through periods of very high inflation. Stopping high inflation - or avoiding it in the few cases where it has not been a central feature of the process - has been a high priority. At the same time, the surge in the overall price level in most of these countries was associated with price liberalization Prices of goods moved rapidly towards international levels. "The freeing of price-setting allowed the forces of arbitrage to drive prices of tradable goods towards those prevailing on world markets. In contrast, service prices first lagged but then started to catch up with the prices of goods, as services became more

[1] McCallum (1995) used a unique data set of trade flows between Canadian provinces and between Canadian provinces and U.S. states to estimate a gravity model of trade, whereby trade volumes depend on the economic size of each region and the distance between them. The model showed that once the size of a U.S. state economy and the distance between it and a Canadian province were taken into account, trade between it and a province was only about 5 percent of the trade that would have occurred if the state had been part of Canada. He concluded that if even the "relatively innocuous Canada-U.S. border continues to have a decisive effect on continental trade patterns … national borders in general continue to matter"(p. 622).

commercialized" (Koen and De Masi, 1997, p. 9). Ukrainian national price changes are in line with these facts highlighted by Koen and De Masi concerning price level convergence in transition economies. Thus, during the period from December 1990 to October 2002, the overall CPI index increased 339 000 times, prices for food increased 310 000 times, for non-food – 159 000 times, while prices for services increased 25 665 000 times (Revenko, 2002, p. 8).

At the same time, Koen and De Masi state that over time, prices and inflation rates have converged across regions within countries, as well as across countries. Since Ukraine follows the common price level adjustment pattern of other transition economies, prices within the country should also converge. The reason to expect at least some regional Ukrainian price convergence is that progress toward a single market, including the already-completed trade liberalization, should narrow differences in prices across oblasts. It means that if prices were initially different across Ukrainian oblasts, convergence to a common level of prices implies higher inflation in oblasts with lower prices (Rogers, 2000)[1]. Therefore, studying factors that cause persistent deviations from the law of one price can help in explaining regional inflation differences.

Most research that tried to identify the causes of inflation in Ukraine investigates only the monetary and institutional sides of inflation. At the same time, Arsenyuk (2002) points out that the Ukrainian deflationary process in 2002 is not linked to monetary policy, but rather is a fault of institutional factors: budget, fiscal and external economic policy. In addition, at the time of writing this thesis I was unable to find any research that tried to analyze regional price differences in Ukraine and their possible link to regional inflation differences.

[1] In theory it might be the case that regions that have higher initial price levels may have deflation, i.e. convergence to the lower price levels may occur instead. However, such a scenario is unlikely in the world of downward price rigidity.

This study tries to answer the question of whether price level in different oblast is the same. Differences between individual regions are found to reach 30%. How the speed of adjustment to the average Ukrainian level differ across different goods? How the average speed of adjustment differ across Ukrainian oblasts?

The rest of the paper proceeds as follows: chapter 2 discusses theoretical background of the law of one price. Empirical findings of previous regional studies are discussed in chapter 3. Panel unit root tests and TAR model estimation procedures are outlined in chapter 4. Data description and preliminary analysis of aggregated price indexes is done in chapter 5. Results of application of panel unit root tests and Band-TAR model are given in chapter.6.

Chapter 2

THEORETICAL FRAMEWORK

As already stated in the introduction, the law of one price, as an extension of 'purchasing power parity' (PPP) for separate goods is an empirical proposition that in competitive markets free of transportation costs and official barriers to trade (such as tariffs), identical goods sold in different locations must sell for the same price when their prices are expressed in terms of the same currency (Krugman and Obstfeld, p. 395).

The difference between PPP and LOP is that the law of one price applies to individual commodities, while PPP applies to some general price level, which is a composite of the prices of all the commodities that enter into the reference basket. If LOP holds true for every commodity PPP must hold automatically. On the other hand, validity of PPP does not require the law of one price to hold exactly. Therefore, the law of one price is more restrictive than PPP.

The reason why LOP must hold is arbitrage. It means that if in region r there is a demand shock that causes a price in this location to rise, there will be a possibility to buy a good at lower price in another location s and sell it for higher price in location r ; therefore, the effect of higher demand will be canceled out by higher supply. In a perfectly integrated economy, were there are no obstacles for the movement of goods between different regions and all markets are competitive, arbitrage is costless (there are no transportation costs), the supply of good i will be perfectly elastic in each location. This is because any increase in demand in location r will be satisfied by an instant inflow of good i from location s (s can be considered as a rest of the country), and a decrease in demand will cause an instant outflow to s. Fig. 1 illustrates different equilibria in a perfectly integrated market in region r.

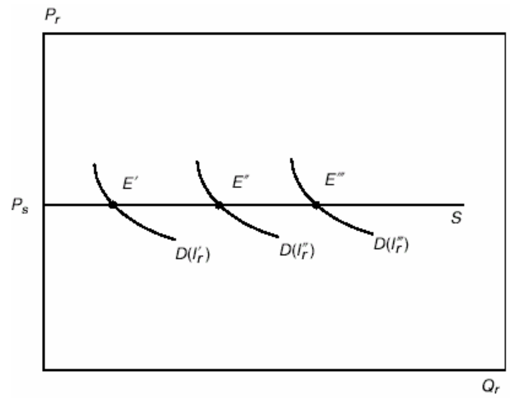

Fig. 1. Market equilibria in region r in perfectly integrated economy with zero transaction costs (C_{rs}=0). D(I_r) is a demand in region r that is dependent on regional income I_r. *Sourse*: Glushchenko (2002).

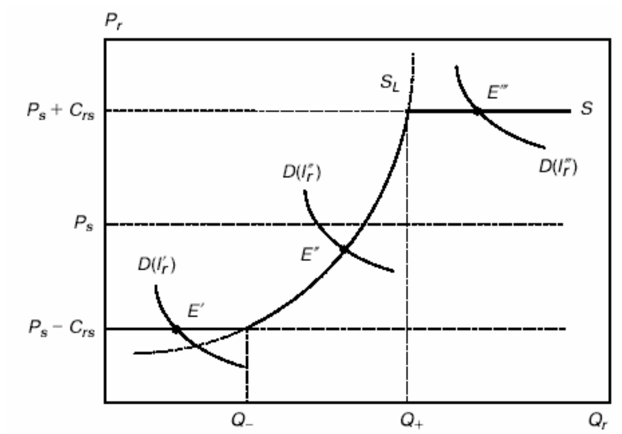

Fig. 2. Market equilibria in region r when transaction costs are nonzero (C_{rs} >0). *Sourse:* Glushchenko (2002).

With nonzero arbitrage costs (transportation costs, trade barriers) C_{rs} >0. However, arbitrage will not be able to equalize prices in regions r and s because arbitrage will occur only if the price difference exceeds transaction costs C_{rs}, otherwise it will be unprofitable. Fig. 2 illustrates the situation.

As it can be seen from the graph, supply of good i in location r is perfectly elastic in section $Q_r \geq Q_+$ when price exceeds ($P_{i,s}$+ C_{rs})because any further effect of increased demand will be cancelled out by increase in deliveries from s that are now profitable. Supply is perfectly elastic in section

$Q_r \leq Q_-$ because any excess of supply can now be exported to s and sold at $P_{i,s}$. However, if demand in location r is such that price of good i $P_{i,r} \in [P_{i,s}-C_{rs}, P_{i,s}+C_{rs}]$, then arbitrage is unprofitable and persistent differences between prices $P_{i,r}$ and $P_{i,s}$ can be maintained.

It is possible not only that demand changes but also the supply of good i changes as well. For example, wages and as a result, production costs rise in region r. In this case bounds Q_- and Q_+ may change, as well as the shape of the supply curve in the interval $[Q_-, Q_+]$. However, this does not change the conclusions about the behavior of the difference in prices.

There are two types of LOP. The first is the *absolute law of one price*. Let $P_{i,r,t}$ and $P_{i,s,t}$ be prices of good *i* at time *t* in locations *r* and *s* respectively. Then, according to the absolute law of one price:

$$\frac{P_{i,r,t}}{P_{i,s,t}} = 1 \qquad (1)$$

This definition of LOP implies that when trade is open and costless, identical goods must trade at the same price regardless where they are sold.

So far, I have focused on absolute LOP. However, there is another version of the LOP, the *relative law of one price*:

$$\frac{P_{i,r,t} / P_{i,r,t-1}}{P_{i,s,t} / P_{i,s,t-1}} = 1 \qquad (2)$$

The relative LOP requires only that the rate of the price change of good *i* in both locations be the same. In this sense it is less restrictive. However, the absolute LOP is less likely to hold in practice because of transportation costs and trade barriers.

There are some difficulties associated with the relative LOP. The relative LOP may hold because:

a) $\begin{cases} P_{i,r,t} = P_{i,s,t} \\ P_{i,r,t-1} = P_{i,s,t-1} \end{cases}$ (absolute prices are equal in all periods: the law of one price holds in all periods)

b) $\dfrac{P_{i,r,t}}{P_{i,s,t}} = \dfrac{P_{i,r,t-1}}{P_{i,s,t-1}}$ (relative prices are equal in all periods: the price in location r is proportional to the price in location s in all periods)

c) $\dfrac{P_{i,r,t}}{P_{i,r,t-1}} = \dfrac{P_{i,s,t}}{P_{i,s,t-1}}$ (changes of prices in both locations are equal)

Therefore, if the available data is in the form of price indexes ($P_{i,r,t}/P_{i,r,t-1}$), it will be impossible to distinguish which relationship holds exactly. In addition, price indexes are not appropriate for testing the absolute LOP (as it will be explained in detail in the data analysis section).

Chapter 3

LITERATURE REVIEW

Researchers have suggested a number of explanations for incomplete relative price adjustment. Factors that can be used to explain deviations from the law of one price between regions within a country are:

i) **transportation costs associated with moving goods from one region to another.** Parsley and Wei (1995) investigate the influence of transportation costs on relative price differences between 48 cities of the United States using a panel of 51 prices. They approximate transportation costs by distance between the cities. Regressing the intercity log of relative prices $Q_{i,rs,t}$ on the log of distance and the log of distance squared (in order to capture nonlinear relationship) they conclude that price differences are bigger for cities that are further apart. Moreover, distance also influences the speed of convergence: convergence rates are slower for cities further apart. However, this influence is small. The limitation of this approach is that the implicit assumption that transportation costs are constant in time is made, whereas it is clear that changes in prices for gasoline can significantly influence transaction costs.

ii) **sticky nominal price level adjustment arising from imperfectly competitive product markets where price changes are costly.** Engel and Rogers (1999) using price indexes data for 29 U.S. cities and for 43 different goods construct a measure of relative price volatility. Surprisingly, they find that variability is larger for traded goods than for non-traded goods and goods with a large nominal price variance show large deviations from the LOP, irrespective of the distance between locations. In fact, deviations from the LOP for non-traded goods were found to be smaller than for traded goods, which the authors attribute to greater nominal price stickiness for non-traded goods (deviations are small because there is not much variation).

iii) **the degree of tradability of goods.** In general, in order to distinguish between tradable and nontradable goods, researchers divide goods into two subgroups. Crudely, all manufactured goods are considered to be tradable, while services are referred to as nontradables. Parsley and Wei (1996) divide goods into these two subcategories and find half-lives for each good. Then, median half-lives for each group are compared. They find that the speed of convergence for services is three times lower than that for tradable goods. Glushchenko (2001) uses the overall regional CPI index for 7 regions of Russia and considers separately its subcategories: food, manufactured goods and services in order to find differences in the speed of convergence between different groups of goods.

In addition to the factors mentioned above Glushchenko (2002) also considers some additional factors specific for transitional countries that influence the price convergence process. Among them:

- **Shipping conditions**. In addition to distance, two variables are adopted to proxy shipping costs: the quality of the region's transportation infrastructure and regional freight tariffs.

- **State intervention in economy.** Two variables represent this factor: price regulations and subsidies. The former is the proportion of goods and services with regulated prices in the region. The latter are production subsidies as a proportion of the regional budget expenditures.

- **Shuttle trade** (small-scale cross-border informal trade). A quantitative evaluation of the phenomenon by region defined as shuttle trade in the region normalized to the average Russian one.

- **Organized crime.** Two proxies of organized crime are used: the number of registered crimes per 10000 of the population and the economic power of crime determined by the proportion of the regional economy controlled by criminal groups.

In line with the possible factors that cause relative prices to be different across regions of one country, empirical work have found several regularities that persist in different samples. Among them are:

i) **adjustment is faster when shocks are large (a nonlinear relationship between speed of convergence and the size of deviation).** Cecchetti, Mark and Sonora (2000) calculate the log of relative prices $q_{i,rs,t}$ and divide the sample into two parts: one part of the sample that has larger $q_{i,rs,t}$ and another one that has smaller $q_{i,rs,t}$. They find the speed of convergence for each subsample and use a Wald test to conclude that the two estimates are significantly different and the speed of convergence for large deviations from the LOP (with bigger $q_{i,rs,t}$) is higher. The second methodology used to capture nonlinear relationship between the speed of convergence and the magnitude of deviation from the LOP is adopted by Obstfeld and Taylor (1997) and O'Connell and Wei (1997). They use TAR model (threshold autoregressive model) that takes into account nonzero transaction costs. According to TAR model, arbitrage takes place only when its gain exceeds transaction costs. If the price gap is inside the band, arbitrage does not affect the ratio of prices.

ii) **nontradable goods prices converge more slowly than those of traded goods.** In addition to discussion of the fact that degree of good's tradability influences its half-life, more evidence can be added by considering the work of Koen and De Masi (1997). This identifies the several well-known features of inflationary processes in Central and Eastern Europe, the Baltics, Russia, and other countries of the former Soviet Union during the years of transition. Koen and De Masi (1997) point out that "service prices first lagged but then started to catch up with the prices of goods, as services became increasingly commercialized." The reason for such behavior of service prices is administrative controls and slower productivity growth in nontradables sector.

iii) **speed of convergence depends on the frequency of the data used.** Taylor (2000) provides extensive explanation of why, and in which cases, temporal aggregation may cause bias in estimates of the speed of convergence. He proves that consensus estimates of half-lives in the range 3-5 years may be a result of temporal aggregation since, typically, PPP and LOP have been tested with aggregate data at annual, quarterly, and monthly frequencies. "Sampling the data at low frequencies will never allow one to identify a high frequency adjustment process. Instead, a large bias could be introduced towards the finding of a long half-life, and the bias grows larger the greater the degree of temporal aggregation."(Taylor, 2000)

iv) **relative price changes and inflation are positively correlated.** A variety of theories can explain this correlation, but the existing evidence is inconclusive concerning the causal mechanism that generates the observed facts. Using panel data for the period 1954-1986 for 19 U.S cities and for 14 categories of goods and services Debelle and Lamont (1996) find that U.S. cities that have inflation above the national average level in a given year also have a higher degree of relative price variability. This relationship holds for different categories of goods and services, for different subperiods, and controlling for fixed city and year effects.

Kaplow (1995) points out the consequences of violation of the law of one price with application to system of taxation. The problem is that national income taxes do not take into account regional variation in price levels (cost-of-living variations), nor do major welfare programs. However, it is generally accepted that changes in the cost of living over time should be reflected in the tax system and the rate structure of income tax should be indexed by regional inflation levels. The article undertakes a preliminary conceptual investigation of regional cost-of-living adjustments in the tax system and examines whether adjustments are efficient. To explore the issue, a benchmark that preserves equal utility between regions was offered.

The following discussion in this thesis will parallel the discussion in Glushchenko (2001) and Glushchenko (2002). Glushchenko (2001) uses CPI indexes for 7 regions of Russia for the period 1992-1998 to determine whether price liberalization in 1992 caused price convergence between Russian regions and to determine the extent of the influence of factors blocking price equalization across regions. Aggregated data is used for the analysis: the overall consumer price level and its subcategories: food, manufactured goods and services. A TAR model is used to find estimates of the speed of convergence for each region to the average Russian level. It is found that the overall price level tends to converge to the average national level in about half of the regions during the whole 1992-1998 period. However, the magnitude of threshold levels in some cases is abnormally high and cannot be explained solely by transportation costs.

Glushchenko (2002) extends these findings, arguing that the relationship between price differences and per capita demand differences (regional income differences are used as proxy) across regions can be used as a cross sectional test of the law of one price. The analysis covers 74 out of 89 Russian regions. In general, a strong relationship between regional incomes and prices is found, which means a poor integration of the consumer market of the country. To evaluate the importance of factors that hinder regional price convergence such factors as the quality of the transportation structure, subsidies, regulations and organized crime variables are incorporated into the analysis.

There are two papers that studied price convergence in Ukraine. The first paper (Cushman, MacDonald, Samborsky, 2001) uses semi-monthly prices for 5 commodities in Kyiv relative to U.S. counterparts during Ukraine's early transition period of 1991-1992. Panel unit root tests are applied to the resulting commodity real exchange rates. Although the LOP did not hold during the period, the commodity real exchange rates appear to have possessed deterministic trends that were in the direction of closing the initial considerable price gap.

The other paper (Conway, 1999) studies the evolution of daily prices of three commodities at four market locations (bazzars) in Kyiv. The interconnected effects of increased market orientation and of privatization are separated and evaluated through the use of an error-correction statistical methodology. The data cover the period from April 12, 1993 to December 31, 1996. Conway found significant evidence of price convergence due to arbitrage by buyers and sellers at these markets *within Kyiv*. A secular trend toward greater market integration over time, and a one-off reduction in price differentials associated with the privatization of former state shops are also found.

Empirical findings of research that investigated the issue of regional price convergence can be summarized in the following table:

Table 1. Summary of previous empirical findings.

Name	**Year**	**Sample**	**Period**
Parsley D., Shang-Jin Wei	July 1996	51 prices from 48 cities of U.S. (quarterly)	1975:1-1992:4
Engel Ch., Rogers Jh.	July 1999	43 prices from 29 cities of U.S. (monthly)	1986:12-1996:6
Alberola E., Marques J.	1999	Overall CPI index for 50 provinces of Spain (quarterly)	1961:1-1998:1
Ceccetti S., Mark N., Sonora R.	May 2000	CPI index for 19 cities of U.S. (annual)	1918-1995
Glushchenko K.	2001	CPI and its 3 subcategories for 7 regions of Russia (monthly)	1992:1-1998:12
Cushman D., MacDonald R., Samborsky M.	May 2001	5 commodity prices in Kiev relative to U.S. counterparts (semi-monthly)	1991:6-1992:4
Chaudhuri K.	July 2001	CPI indexes of 8 goods for 7 cities of Australia (quarterly)	1972:4-1999:1

Chapter 4

METHODOLOGY

Most work that tries to determine whether the law of one price holds or not, follows a standard procedure.

Logarithm of relative prices.

At the first step, **relative prices are obtained.** The fact that by the law of one price, perfect arbitrage will force prices of all tradable goods to be the same is employed. Thus, the LOP benchmark for

$$\mathbf{q_{i,rs,t} = \ln(P_{i,r,t}/P_{i,s,t})}$$

is 0[1] (where i is an index of good; r and s are the regions of interest and t is month in a particular year). However, in practice log of price ratio $q_{i,rs,t}$ can be different from 0 due to the cost of arbitrage between the two regions. As a result, $q_{i,rs,t} = \varepsilon_{i,rs,t}$. If the law of one price holds it means that shocks, that force $q_{i,rs,t}$ not to be equal to 0, do not have permanent effect on relative prices; $\varepsilon_{i,rs,t}$ is then stationary noise with mean 0 (prices are cointegrated). If LOP does not hold then $\varepsilon_{i,rs,t}$ is nonstationary.

However, in practice $\varepsilon_{i,rs,t}$ may not have mean equal to 0 because of random shocks, transaction costs etc. In order to eliminate cross-sectional specific effects, series $q_{i,rs,t}$ are demeaned before the analysis is done.

In the standard model $q_{i,rs,t}$ is assumed to follow AR(1) process:

$$\mathbf{\Delta q_{i,rs,t} = \lambda_{i,rs} * q_{i,rs,t-1} + e_{i,rs,t}} \qquad \text{AR(1)}^2$$

1 It follows from the fact that by the law of one price $P_{i,r,t}/P_{i,s,t}=1$

2 There is no intercept in the specification because series $q_{i,rs,t}$ are already demeaned.

To check for stationarity of $q_{i,rs,t}$ univariate augmented Dickey-Fuller test is used[1]. H_0 is the unit root hypothesis: $\lambda_{ij,r} = 0$. The alternative hypothesis is stationary $q_{i,rs,t}$: $\lambda_{i,rs} < 0$ (prices converge). $\lambda_{i,rs}$ is speed of convergence, the speed with which market forces bring prices back to an equilibrium. The convergence speed is usually interpreted as a measure of market's integration or the efficiency of arbitrage between spatially separated locations, and is expected to depend on the good or composite goods under consideration, the nature of transaction or transportation costs for these goods etc. The derivative indicator that captures the price gap between regions is called the "half-life time of price gap" and is calculated as:

$$\mathbf{T_{i,rs} = \ln 0.5/\ln(1+ \lambda_{i,rs})}$$

It determines the number of time periods it takes to reduce by half the magnitude of a deviation from the equilibrium caused by an individual shock (price gap between regions r and s).

Panel unit root tests.

However, "in finite samples, univariate augmented Dickey-Fuller test inevitably has limited power against alternative hypothesis with highly persistent deviation from equilibrium" (Levin and Lin, 1993). Thus, it is very difficult to reject unit root with ADF when in fact hypothesis of unit root is false. The natural response to the slow mean reversion properties of the observed movements in the relative prices is to increase the span of the data in both time and individuals dimensions. Panel unit root tests allow one to overcome some of the problems associated with univariate unit root tests. Their main advantage is increasing sample size by pooling the data.

Most studies that are written during the period from 1970s to early 1990s and make pairwise comparisons of countries used univariate methods to

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[1] Univariate augmented Dickey-Fuller test will not be used in this study.

check for stationarity. They typically did not reject the hypothesis that $\varepsilon_{i,rs,t}$ is nonstatinary[1]. This result implies that inflation differentials among countries, when measured in common currency, can persist indefinitely or, in other words, that the price level in one country can deviate by an arbitrary large amount from that of another.

In contrast, studies that were written in late 1990s and used multivariate tests that combine numerous countries in panel unit-root testing procedures were able to reject the unit-root hypothesis, which implies that relative prices revert to a common mean. However, the rate at which this convergence occurs is slow. Deviations appear to damp out at approximately 15% per year, that is, the consensus among these studies on the half-life estimates is in the range of 3-5 years.

The most popular panel unit root test are LL test (Levin and Lin, 1993) and IPS test (Im, Pesaran, Shin, 1997). They are generally used for panels of moderate size (i.e. between 10 and 250 cross-sections, with 25 to 250 time series observations per one cross-section).

Both panel unit root tests assume that data follow the stochastic process:

$$\Delta q_{i,rst} = \alpha_{i,rs} + \theta_{i,t} + \beta_{i,rs} q_{i,rs,t-1} + \sum_{j=1}^{k} \gamma_{i,rs,j} \Delta q_{i,rs,t-j} + \varepsilon_{i,rs,t}$$

where $q_{i,rst}$ is the log of relative price of regions r and s for good i at time t, $\alpha_{i,rs}$ is a cross-section (region pair) specific constant to control for non-time-dependent heterogeneity across cross-sections, and $\theta_{i,t}$ is a common time effect. The $\gamma_{i,rs,j}$ are lag coefficients in the process characterizing $q_{i,rs,t}$, $\beta_{i,rs}$ is the coefficient that is used to compute half-life of a shock.

$$T_{i,rs} = \ln 0.5 / \ln(1 + \beta_{i,rs})$$

[1] Good survey of literature up through the early 1990s can be found in Rogoff (1996).

$\alpha_{i,rs}$ allows for cross-sectional heterogeneity, such as different income levels and sales taxes that can lead to permanent differences in prices in different oblasts. $\theta_{i,t}$ is a common time variable that captures common macroeconomic shocks that influence all cross-sections simultaneously and is equivalent to time trend[1]. $\varepsilon_{i,rs,t}$ are assumed to be independently distributed among cross-sectional units. Researchers are interested in coefficients $\beta_{i,rs}$ which should be negative. The closer are the estimates to 0, the longer the estimated half-life of a shock and the more likely it is that prices are nonstationary.

LL and IPS tests differ in their treatment of $\beta_{i,rs}$. Levin and Lin (1993) assume that $\beta_{i,rs} = \beta$ for all pairs r and s for every good i. This means that the coefficient of the lagged dependent variable is assumed to be homogeneous across all cross-sectional units of the panel. The null hypothesis in both tests is the same, i.e. that all cross-sectional units are nonstationary, H_0: $\beta_{i,rs} = \beta = 0$. However, these tests have different alternative hypotheses. In the LL test the alternative hypothesis, H_1: $\beta_{i,rs} = \beta<0$, i.e. all cross-sections are stationary. In contrast, IPS allows for heterogeneous coefficient $\beta_{i,rs}$ in cross-sections but the alternative hypothesis of IPS test H_1: $\beta_{i,rs}<0$ is that at least one of the individual series in the panel is stationary (see Appendix A1 and A2 for detailed procedures of estimation).

The Levin and Lin test is restrictive in the sense that it requires $\beta_{i,rs}$ to be homogeneous across cross-sections. On the one hand, it was found in several cases that IPS test has more power than LL test. On the other hand, LL test is more convenient in that it gives one estimate of $\beta_{i,rs}$, whereas IPS does not provide a single estimate of the speed of convergence. In addition,

[1] Trend will not be included into the specification of panel unit root tests in the analysis of regional Ukrainian retail prices.

even if IPS test rejects H_0 of a unit root, it does not say how many and which cross-sections in the panel are stationary.

Test statistics of both tests have nonstandard distributions and bootstrapping should be done to find p-value of calculated t-statistic. However, Levin and Lin (1993) show that as both the cross-section and time series dimensions of the panel grow large, the panel unit root statistic has a limiting normal distribution. Levin and Lin (1993) therefore provide correction and standardization factors required for the unit root estimators to have a normal distribution in the limit. Critical values of IPS statistic are derived by Monte Carlo simulation and tabulated in the original paper of Im, Pesaran, Shin (1997)[1].

Threshold autoregressive model (TAR).

Taking into account nonzero transaction costs. As it has been said above, a simple AR model is usually used to study the potential convergence of prices toward the law of one price. Convergence speed $\lambda_{i,rs}$ is then interpreted to measure market integration or efficiency of arbitrage. However, if there are market frictions like transportation costs, tariffs, non-tariff barriers, menu costs or pricing to market behavior then there can be deviations from the law of one price without arbitrage taking place. These frictions create a band of no convergence (Obstfeld and Taylor, 1997).

A TAR model is used to capture this effect. According to this model, arbitrage takes place only when its gain exceeds transaction costs. If the price gap is inside the band, arbitrage does not affect the ratio of prices. In the TAR presentation, the observations are split into two regimes, the

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[1] LL and IPS tests were performed using Stata module written by Fabian Bornhorst (European University Institute, Italy) and Christopher F Baum (Boston College, USA).

inner regime where there is little or no adjustment and the outer regime where large deviations make arbitrage profitable and there is convergence to the law of one price.

There are basically two different TAR models to choose from: the EQ-TAR and the Band-TAR.

$$\Delta q_{i,rs,t} = \begin{cases} \lambda^{out}(q_{i,rs,t-1} - c) + e^{out}_t & \text{if } q_{i,rs,t-1} > c \\ \lambda^{in} q_{i,rs,t-1} + e^{in}_t & \text{if } c \leq q_{i,rs,t-1} \geq -c \\ \lambda^{out}(q_{i,rs,t-1} + c) + e^{out}_t & \text{if } -c > q_{i,rs,t-1} \end{cases} \quad \textbf{Band-TAR (1;2;1)}$$ [1]

$$\Delta q_{i,rs,t} = \begin{cases} \lambda^{out} q_{i,rs,t-1} + e^{out}_t & \text{if } q_{i,rs,t-1} > c \\ \lambda^{in} q_{i,rs,t-1} + e^{in}_t & \text{if } c \leq q_{i,rs,t-1} \geq -c \\ \lambda^{out} q_{i,rs,t-1} + e^{out}_t & \text{if } -c > q_{i,rs,t-1} \end{cases} \quad \textbf{EQ-TAR (1;2;1)}$$

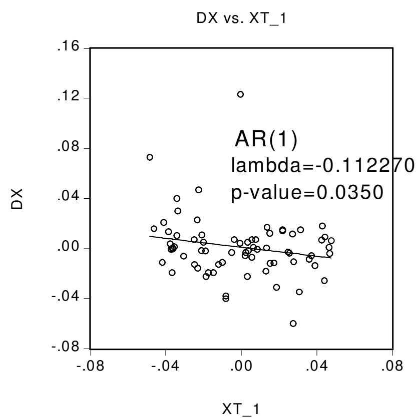
In both models $q_{i,rs,t}$ are the series of interest (log of relative prices of good i in regions r and s) with λ^{in} and λ^{out} being the adjustment coefficients, c is the threshold that separates two regimes, e^{out}_t and e^{in}_t are the noise. EQ-TAR model exhibits mean reversion toward the mean of the series, while Band-TAR model represents process that reverts to the edge of the threshold. The process is stationary overall if the outer band dynamics are stationary: the process always reverts to the inner band in this case.

The equilibrium in Band-TAR model is achieved whenever $q_{i,rs,t}$ is inside the band, i.e. $c \leq q_{i,rs,t} \geq -c$. In contrast to panel unit root tests equilibrium will hold at any point inside the band [-c;+c] and not just at point 0. Since there is no arbitrage inside the bounds [-c;+c], $q_{i,rs,t}$ may follow random walk, drift or stationary process. Therefore, in many studies λ^{in} is restricted to be equal to 0 or not reported. Both coefficients λ^{in} and λ^{out} should be negative but, typically, λ^{in} coefficients are

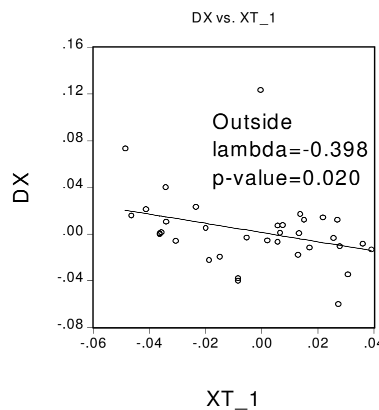
[1] Note that series $q_{i,rs,t}$ are already demeaned.

insignificant and $\lambda^{in} > \lambda^{out}$, implying faster convergence for relatively large price differences outside the band [-c;+c] (Obstfeld and Taylor, 1997).

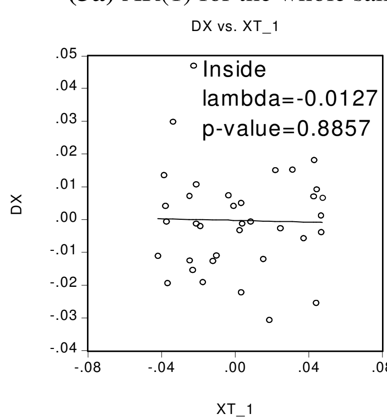
Testing for the law of one price by panel unit root tests that assume AR(p) process instead of taking into account thresholds may lead to a biased estimate of λ if there are transaction costs.



(3a) AR(1) for the whole sample



(3b) Outside-the-band regression



(3c) Inside-the-band regression

Fig. 3. Biased estimate of AR(1) versus TAR model.
Graphs are drawn for relative prices or butter of Dnipropetrovska oblast (region 4) to average Ukrainian price level. Variable ($q_{i,,rs,t-1}$ - c) is on the horizontal axis, variable $\Delta q_{i,rs,t}$ is on the vertical axis.

The problem with the usual AR(1) model is that it does not distinguish between observations outside and inside the band and ignores the threshold. Figure 3 shows what happens when AR(p) model that is applied in panel unit root tests pools the data as if they are a uniform process.

As it can be seen from the graphs, $\lambda^{in} > \lambda^{AR(1)} > \lambda^{out}$, which means that the speed of convergence for observations outside the band is faster than for the points inside the interval [-c;+c], and a slope of the AR(1) model that pools the data is in between λ^{in} and λ^{out}. λ^{in} is insignificant.

The models proposed above are from a large family of TAR (p,k,d) models, which may be characterized by arbitrary autoregressive length p, an arbitrary number of thresholds k, and an arbitrary delay parameter d. The thresholds in specifications shown above are symmetrically placed around mean of the series with the same adjustment coefficient λ^{out} in both out-of-the-band parts of the model. It means that the speed of convergence is the same regardless whether the difference in prices in regions r and s is above the upper threshold or below the lower threshold. The reason to assume symmetric thresholds is that value of a threshold is assumed to represent transaction costs, for example, distance or shipping costs (Johansson, 2001). Since there is no reason to assume that shipping costs, for example, from Kyiv to Donetsk are different from shipping costs from Donetsk to Kyiv, symmetric threshold model is chosen for the analysis of regional price differences across Ukrainian oblasts.

There are several reasons to use Band-TAR model rather than EQ-TAR. Firstly, if it is true that there are transactions costs and arbitrage is profitable only if deviations from the law of one price are big enough, then there should be some interval in which convergence to 0 does not occur, i.e. convergence occurs not to 0 but rather to a band. This is exactly the effect that Band-TAR model captures. Secondly, Band-TAR model is continuous at the threshold. It means that specification tests of Band-TAR model against usual AR(1) model is also a test for $\lambda^{in} = \lambda^{out}$, i.e. the test of whether the convergence speeds are the same in two regimes (Obstfeld and Taylor, 1997). Therefore, Band-TAR model is used for the analysis of regional price differences in Ukraine.

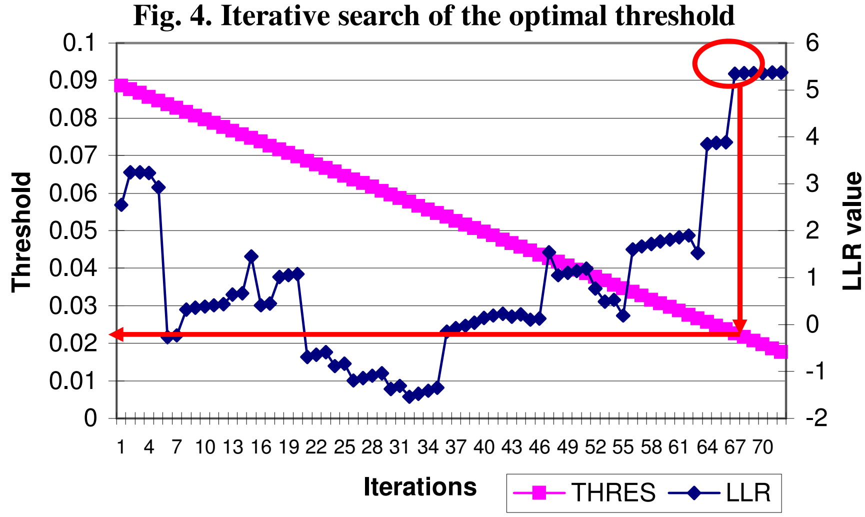
In order to test for a particular specification of TAR as alternative against AR(p) null, LLR test is performed:

$$\mathbf{LLR=2*(L_{TAR} - L_{AR}),}$$

where L_{TAR} is the value of the log likelihood function of Band-TAR model and L_{AR} is the log-likelihood function of AR(1) specification.

Band-TAR model has three unknown parameters: λ^{in} , λ^{out} , and a threshold parameter c. The task is to find MLE or OLS estimates of λ^{in} and λ^{out} for a given threshold c from partitioned samples with observations $q_{i,rs,t-1}$ inside and outside the threshold. The optimal threshold c is found by a stepwise search of c that maximizes log likelihood ratio LLR (see Appendix B for detailed procedure of TAR model estimation).

Fig. 4. Iterative search of the optimal threshold



Since the threshold c is not identified under the null hypothesis of AR(1) model LLR test has a nonstandard distribution and Monte Carlo simulation should be used to find critical values of this distribution. “Half life” of the price gap is calculated by the same formula as in panel unit root tests.

Chapter 5

DATA

Data description.

Regional price and inflation rate differences in Ukraine are not well documented. The monthly oblasts' consumer price index (CPI) was calculated by Derzhcomstat starting only from January 2001. Before this date only the national CPI is available. As a substitute for consumer prices, the retail price index (RPI) and nominal retail prices can be used. The regional RPI is calculated by Derzhcomstat starting from January 1995. Besides RPI, the only indicator that can be used to track oblasts' differences in the cost-of-living is "dekadka", a cost of fixed basket of 22 food products[2] that have the largest shares in household consumption. Derzhcomstat calculated "dekadka" from January 1992 to December 1998. However, the data for "dekadka" used in this study are for the period from 1995:01 to 1998:12 (48 time series observations for each cross-section). The data for national and regional monthly RPI are for the period from January 1998 to December 2002.

Nominal retail prices of 30 goods and services are used in the analysis. Prices are divided into subgroups of food, nonfood and services following official classification of Derzhcomstat. The analysis of regional price differences in Ukraine includes 17 food, 9 nonfood goods, and 4 services.[2] Goods and services were included into analysis by the principle of the largest share in the household consumption and availability of the data. The data for

1Basket of 22 food commodities include: flour, bread, noodles, semolina, beef, pork, chicken-meat, sausage, milk, cheese (hard and soft), sour cream, butter, eggs, sugar, oil, potato, cabbage, carrot, beet, onion, apples. 17 out of 22 of these goods are used in the analysis of regional retail prices.

2 See Appendix C , table C1 for detailed description of each item.

food and nonfood commodities are available from January 1997 to June 2002 for 25 oblasts, Kyiv, Sevastopol, and for the whole Ukraine (72 time series observations for each cross-section). The data for services are available only from January 1997 to December 2001 (60 time series observations for each cross-section) because starting from January 2002 Derzhcomstat has changed classification of services and the data for the later period are not comparable with prices before January 2002. The data for Sevastopol are available starting only from January 1999[1]. Nominal retail prices for this city were excluded from panel unit root tests and TAR model analysis to keep the panel balanced. Several missing values of nominal retail prices were extrapolated.

Absolute prices should be converted to the log of relative prices in order to test the law of one price. We choose the average Ukrainian price level for each commodity as the numeraire (O'Connel and Wei, 1997).

Nominal retail prices are collected in 550 urban locations in all 25 oblasts of Ukraine and the cities of Kyiv and Sevastopol. Cities are selected taking into account the size of population; their representativeness in form of social, economic and geographic situation; the presence of a wide range of products. Registration of prices and tariffs is conducted in retail networks (city markets (bazaars) are not included) and for enterprises that offer services that are included in the list. Enterprises are selected by the local statistical offices that take into account household demand on enterprise's products, their retail turnover and regular trade during the long period of time. The actual price of a good or service is registered including taxes (VAT, excise tax and other indirect taxes). Average nominal retail prices that are published in "Average Prices and Tariffs for Consumer Goods and Services" are used to calculate RPI.

The main difference between consumer and retail prices is that consumer prices include observations from retail shops and city markets

[1] Before January 1999 prices of Sevastopol were included into aggregated prices for the whole Crimea.

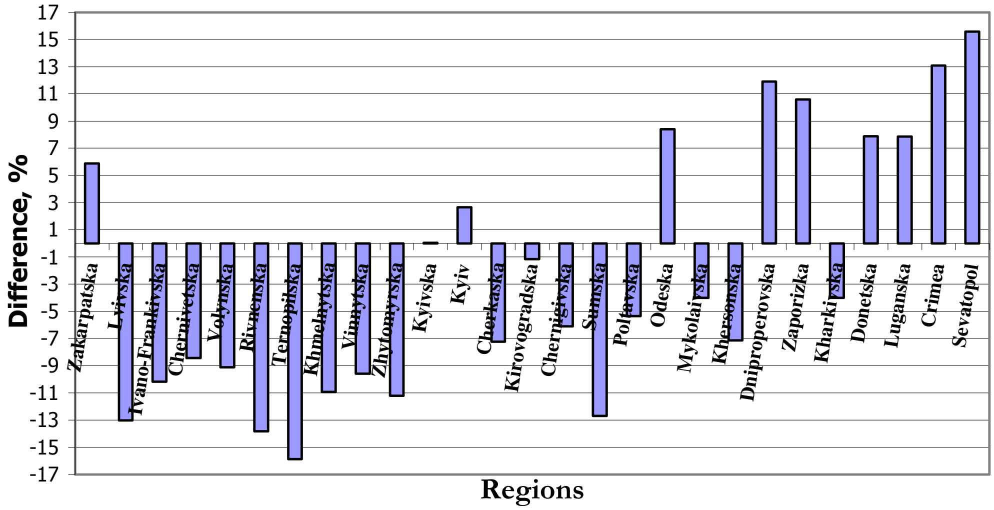
(bazaars), while nominal retail prices and RPI include observations only from retail shops (markets are not included). Usually, prices for food products on city markets are lower and more volatile than in retail shops. The other significant difference between CPI and RPI is the weights that are used to calculate these indexes. CPI uses weights from a structure of household consumption from the Derzhcomstat survey "Household Income and Expenditures in Ukraine" conducted quarterly. CPI weights are quite stable from year to year. The highest weight is given to food (64%) with the most important food products such as: meat (8.8%), bread (7.7%), milk (4.7%), sugar (3.8%), vegetables (3%), potato (2.9%), sausages (2.9%), and eggs (2.5%). Non-food industrial products comprise 14% of the CPI (including clothes - 3,45%). Services comprise 22% of the overall CPI with central heating (3.2%) and electricity (3%) having the highest weights.

In contrast, weights of the RPI are shares of goods and services in retail turnover that change significantly from year to year. Food comprised 68% of retail turnover in 1996 but only 49% in 2002. The distinctive feature of RPI is that it is greatly influenced by a volatile component such as petroleum, that comprised 14% of overall RPI in 2002. It also explains the difference in CPI and RPI dynamics in 2002. CPI in 2002 decreased by 0.6%, while RPI increased by 6.8% due to a 35% rise in the price of petroleum.

Are the regional price levels the same? Analysis of aggregated data.

Convergence of regional prices to the average national level can be investigated by looking at some composite index. The cost of a basket of 22 food commodities, "dekadka", is a good measure of regional price disparities. Figure 5 below shows the differences between the regional cost of 22 food items and average Ukrainian price level in December, 1998.

Fig. 5. The difference between the cost of a basket of 22 food goods in oblast and avearge Ukrainian level (December, 1998)



The differences of price levels in some regions from the average Ukrainian level in 1998 were from -16% to +16% implying that differences between some individual regions were of approximately 30%. Almost all the regions that had price level higher than the average Ukrainian level at the end of 1998 were the eastern and southern regions[1]. Has the situation changed by 2003?

The problem is that "dekadka" which was the only measure of aggregated regional price levels expressed in nominal terms is not available after 1998. Therefore, we have to rely on indexed data. As stated in the data description section, a regional CPI index is available only for the period 2001-2002, while RPI index is available for the period 1998-2002. The raw RPI data is given in the form of monthly indexes. Given the value of the price level in a region in the initial period (December 1997 is taken as t=0), the *absolute* regional price level at time *t* can be calculated by the chain method[2] as

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1 See the map of Ukraine with the list of oblasts in graph 1 in Appendix C.

2 Price indexes are given in the form $P_t/P_{t-1} = J_{r,t}$ rather than $(P_t-P_{t-1})/P_{t-1}$.

$$\mathbf{P'_{r,t}=P'_{r,0}\times J_{r,1}\times J_{r,2}\times J_{r,3}\times\ldots\times J_{r,t-1}\times J_{r,t}=P'_{r,0}\times I_{r,t}}\ .$$

Regional price level relative to national price level can be obtained as:

$$\mathbf{P_{r,s,t} = P'_{r,t}/P'_{s,t} = I_{r,t}/I_{s,t}}$$

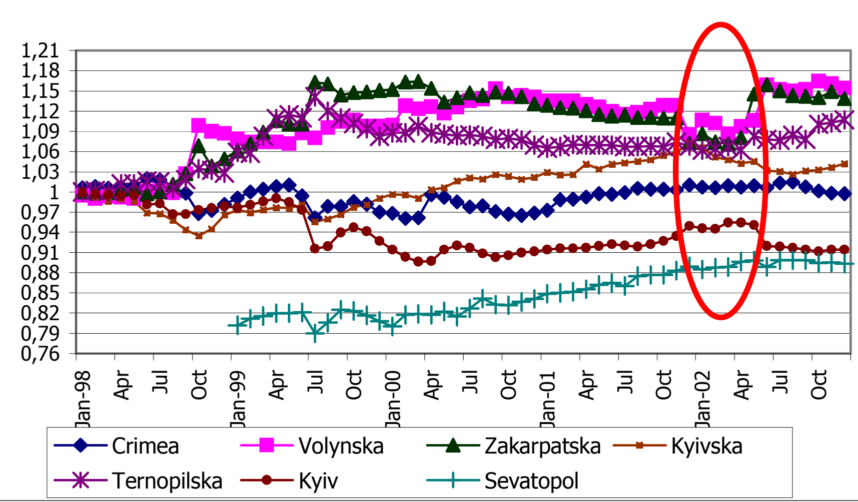
if the initial (at December 1998) regional absolute price levels are assumed to be equal ($P'_{r,0} = P'_{s,0}$). In the analysis of RPI data, regional indexes will be divided only by national index (s=0). Table 2 shows some summary statistics (means and standard deviations over all regions and over all months of a given year) of the relative price levels $P_{r,s,t}$ (where s=0 is a national price level).

Table 2. Summary statistics of relative price levels.

	1998	1999	2000	2001	2002
mean	1.0042	1.0069	1.0209	1.0237	1.0278
std. dev.	0.0225	0.0613	0.0719	0.0652	0.0687

As it can be seen from the table there is no clear evidence that prices tend to equalize since mean and standard deviation do not fall over time (in fact, both mean and standard deviation increase). The dynamics of regional relative price levels for several oblasts are shown in fig. 6.

Fig. 6. Ratio to Ukrainian average level (base is December 1997)

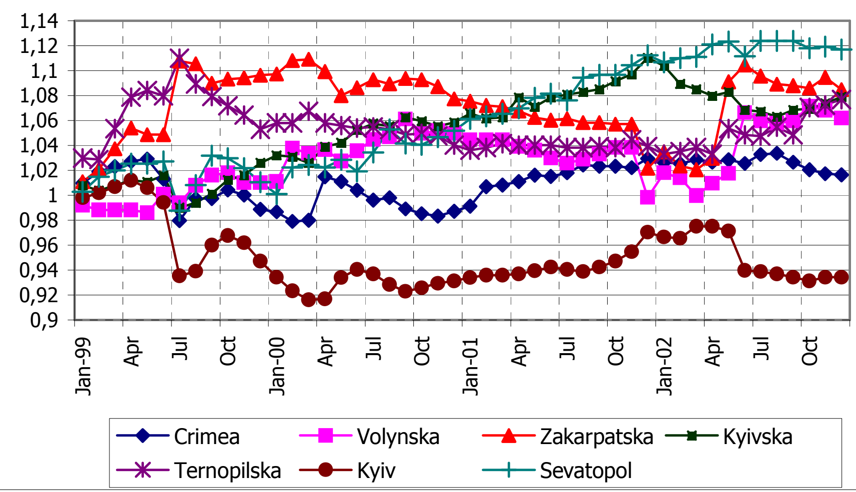


*RPI data for Sevastopol is available only starting from Jan 1999. Before this date Sevastopol was included in calculation of Crimea index.

Fig. 6 shows relative price levels of several regions compared to the average Ukrainian price level. There was some tendency to convergence in Volynska and Zakarpatska oblasts up to the beginning of 2002[1]. However, later on these oblasts show divergence from the national level. Sevastopol, in contrast, shows a stable pattern of convergence to the Ukrainian average price level.

However, the analysis of price convergence that uses price indexes rather than nominal prices should be used with caution because it assumes that all regions had the same initial price level at the base period. Fig. 6 uses December 1997 as a base period and there is no reason to think that the overall price level in all Ukrainian regions was the same at that time. Fig. 7 uses December 1999 as a base period and results are changed substantially.

Fig. 7. Ratio to average Ukrainian level (base is December 1998)



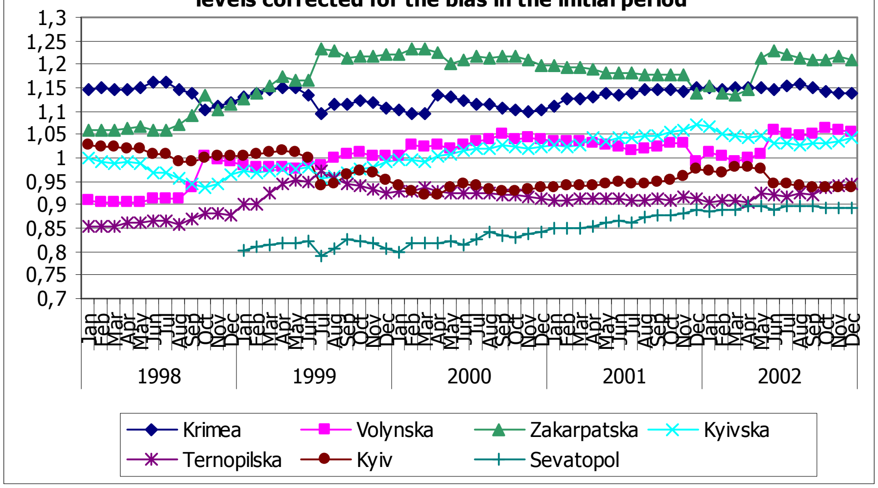
The most dramatic changes have occurred with Sevastopol. Even though the Sevastopol price level was not affected by the change in the base period, its relative index has changed substantially and now it is even larger than the

[1] This pattern is also confirmed by graphs of nominal prices of some goods.

national average. Kyivska oblast also has a relative price level much higher than before.

In order to avoid bias associated with the choice of a base period we could assume that regional differences in the cost of a basket of goods in 1998 could be a proxy for the differences in the overall regional price level. Then, we can multiply RPI index in January, 1999 by the cost of "dekadka" in December, 1998 to account for different price levels in the base period. The results are shown in the fig. 8. below.

Fig. 8. Relative to the average Ukrainian level regional RPI price levels corrected for the bias in the initial period



The pattern of dynamics has changed for some regions, but differences in the regional price levels remain even in 2002.

Thus, the analysis of aggregated indexes indicates that regional price level differences are persistent. It is better to use nominal prices rather than price indexes to test for the absolute law of one price to avoid the bias of the base period.

Chapter 6

EMPIRICAL ANALYSIS

Nominal prices. Panel unit root tests.

There are two issues related to the specification of the panel unit root tests:

1. Trend. The specification of the regression of panel unit root tests in methodology part contains a trend. Both tests were performed with and without deterministic trend in specification. Inclusion of a trend reduced estimated half-lives substantially. When the trend was included estimates fell two times in some cases. This result is consistent with the results of Cecchetti, Mark, and Sonora, 2000. However, as authors point out, detrending of the data is usually used only when the data are in the form of price indices. Due to disadvantages of indexes noted above (due to the problem of the base year, due to aggregation in price indices or because of inclusion of nontradable components), indices may show a stable difference in prices between the two regions. In this case detrending of the log of relative prices should be made. However, detrending of nominal prices may be inconsistent with the theory because there is no economic basis to expect relative prices between Ukrainian oblasts to trend over long horizons. As a result, *only constant is included into the specification* of the panel unit root tests.

2. Number of lags included. Next issue related to the specification of the panel unit root tests is the question of how many lags to include. In case of the analysis of Ukrainian nominal retail prices, it was found that results of the Levin and Lin test (results of IPS test were not sensitive to inclusion of different number of lags) depend on how many lags are included. Table C3 (Appendix C) shows full results of panel unit root tests with different lags included. As it can be seen from the table, the speed of convergence does not depend significantly on the number of lags included in specification. LL test results, however, suggest that rejection of a unit root is less likely when the number of lags included increases. Thus, inclusion of a lag of order greater than 6 usually leads to an acceptance of a unit root hypothesis. There are cases when LL test statistic has high p-value (meaning that all cross-sections in the panel are nonstationary) when 12 lags are included. However, with inclusion of only 4 lags LL test indicate that null hypothesis of nonstationarity can be rejected, i.e. p-value of the speed of convergence fall substantially. Since the data used in the analysis are monthly prices, inclusion of only 4 lags may result in autocorrelation of residuals. Autocorrelation may result in biased results of LL test. Therefore, *inclusion of at least 12 lags is done* to ensure that errors are not autocorrelated.

Results of LL and IPS tests are presented in table C2 (Appendix C). This table shows results of LL and IPS tests when only constant and 12 lags are included into the specification. Table C3 (Appendix C) shows how the speed of convergence and result of the LL test depend on the number of lags included. This table also gives standard deviation of the estimates (in squared brackets) so that the average estimate and corresponding half-life for categories food, nonfood and services could be calculated.

As it can be seen from the table C2, LL test shows that the unit root hypothesis cannot be rejected for all goods and services used in the analysis. P-values of all coefficients are close to 1. It means that the law of one price does not hold for all goods and services that are considered in the analysis. There is one exception though: p-value of the speed of convergence λ for photo services is 0.0294, which means that nonstationarity of relative prices for photo services can be rejected at 5% level of confidence. Moreover, IPS test for this service has also p-value of 0.000. Thus, the fact that prices of photo services in all oblasts of Ukraine converge to the average Ukrainian level is confirmed by both panel unit root tests. However, the speed of convergence is very slow: half-life of deviations is 9 months.

Results for other goods and services, however, indicate that speed of convergence to average Ukrainian level is insignificantly different from 0. This result is confirmed by both tests for all nonfood commodities and for all services with exception of photo services. Still, IPS test, being not so restrictive as LL test, indicate for some food commodities (flour, sugar, oil, potato, rice, beet, cabbage, semolina, and basket of 22 food goods) that some of the regional prices do converge to the average Ukrainian level (H_0 is rejected in favor of H_1). However, IPS test does not show how many and which of the cross-sectional relative prices are stationary. Therefore, we cannot conclude with certainty that the law of one price holds for these goods.

Even if the speed of convergence in LL test is not significantly different from 0 for all goods and services under study (with exception of photo services) , the size of the coefficient may still give an insight into how fast can the regional price levels be expected to converge to the average Ukrainian level.

Calculated estimates from table C2 (Appendix C) show that half-lives for food goods are on average lower than for nonfood and service categories. Thus, almost all estimates of half-lives range from 1 to 4 months. Only mineral water has a half-life of 10 months (this is the only good in food category that is not "agricultural"). In addition, beef and sour cream also have half-lives of approximately 7 months. Cabbage, potato, and beet have the fastest speed of convergence and the lowest half-lives of around 1-2 months in the food category. This can be explained by the fact that markets for these vegetables are very competitive and are not regulated at all.

Table 3 below shows the average speed of convergence[1] and the corresponding half-life for each subcategory of commodities.

Table 3. Aggregated estimates from panel unit root tests.

Category	Speed of convergence	Half-life
Food	-0.193418	3.22
Nonfood	-0.095612	6.9
Services	-0.069887	9.57

It is seen from the table that the speed of convergence of services is three times slower than for food commodities. It takes nearly 10 months on average for services to eliminate half of the magnitude of the price shock. The speed of convergence of nonfood products, in turn, two times slower than the speed of convergence of food goods. These results are in line with the finding of Parsley and Wei (1996) that used quarterly city price indexes for US to perform LL test. They found that half-lives for deviations from parity are three times bigger for services than for perishables (food category in Ukrainian context).

Thus, *we may conclude* that panel unit root tests indicate that the law of one price does not hold in Ukraine (small exception is photo services). It means that according to panel unit root tests regional price disparities do not diminish with time and there is no convergence of regional prices to the average Ukrainian level. At the same time, even insignificant coefficients of the speed of convergence show that the fastest adjustment may occur for food commodities that have competitive markets and are not regulated. The slowest adjustment is expected for services category.

1 Following methodology of Parsley and Wei (1996) median estimates for each category were also calculated but they are very close to the average estimates and, therefore, are not reported.

Taking into account transaction costs. TAR model results.

Results of the estimation of the Band-TAR model are shown in tables C4 and C5 (Appendix C). Band-TAR model was estimated for each cross-section of each commodity. Then, coefficients for 26 cross-sections (25 oblasts and Kyiv) were averaged; their standard deviations, corresponding half-lives, and average threshold levels were calculated for each commodity. Table C4 (Appendix C) shows average coefficients across 26 Ukrainian regions for each commodity. P-values of these coefficients are not reported because all these average coefficients are very significant (variances of such coefficients are very small).

As it can be seen from the table C4 (Appendix C), the speed of convergence from the TAR model for all goods is substantially higher (the coefficient is higher by absolute value) than from the usual AR(1) model that does not distinguish between observations outside and inside the band. Consequently, half-lives from the TAR model are substantially lower than from the AR(1) model. Thus, half-lives from the TAR model range from 0.28 to 1.63 months for food commodities, while the range of half-lives from the AR(1) model for the same commodity group is from 1.37 to 4.83. The same pattern is observed for nonfood and services: even the highest half-life of the TAR model in these groups is lower than the lowest half-life from the AR(1) model. The half-lives for nonfood group from the TAR model range from 1.9 to 2.19 months; for services half-lives are in the range from 1.7 to 2.16 months.

The goodness of fit of the TAR model versus AR(1) model can be checked by LLR test. If the TAR model is better model than AR(1) it is expected that the p-value of the LLR statistic should be as small as possible to be able to reject the null hypothesis of AR(1). The last column of table C4 (Appendix C) shows how many of the cross-sections for each commodity are able to reject AR(1) in favor of the TAR model. However, results suggest that only for several commodities more than 50% of cross-sections reject AR(1).

Only for beet and horilka more than 70% of oblasts confirm that TAR specification is better than simple AR(1) model. However, this result is not surprising given the low power of relatively short univariate time series to distinguish among near unit root alternatives (Obstfeld and Taylor, 1997). Thus, Obstfeld and Taylor using CPI indexes for several US cities with 180 time series observations in each cross-section were able to reject AR(1) in favor of TAR model in only very few cases. In addition, the value of the log likelihood function is not the only criterion for the fit of the TAR model. Coefficients of the TAR model in almost all the cases were more significant than coefficients of AR(1) model. Moreover, there are a lot of cases when according to the LLR statistic TAR model is rejected but coefficient λ^{out} is very significant.

Threshold levels from the TAR model are given in table C4 (Appendix C) as a percentage of the mean of relative prices. Formula of transformation of the nominal threshold into the percentage form is:

$$C=(e^{c}-1)*100$$

where C – is the threshold in percentage form and c is the value of the threshold received from the estimation procedure outlined in Appendix B.

As it can be seen from the table C4 (Appendix C), thresholds are in the range from, approximately, 3% to 15% for food category, from 4% to 12% for nonfood commodities, and from 5% to 7% for services. It is interesting to note that the highest thresholds in the food category, i.e. 13% and 15%, are observed for beet and cabbage. At the same time, these goods also have the lowest half-lives in this food category. The reason is that cabbage and beet have large seasonal price spikes in the first half of the year. These big shocks are eliminated very rapidly. Therefore, the speed of convergence for these goods is very high. However, during the rest of the year prices for these goods are stable and there are only minor fluctuations with persistent regional price differences which explain high threshold level.

Table 4. Aggregated estimates from the Band-TAR model.

Category	LLR	Lambda AR(1)	Lambda TAR	Half-life (LL test)	Half-life (AR1)	Half-life (TAR)	Threshold, %
Basket of 22 food goods	5.94	-0.38	-0.73	1.93	1.47	0.53	2.56
Food	13.64	-0.24	-0.51	3.22	2.48	0.98	5.93
Nonfood	17.36	-0.13	-0.35	6.90	4.83	1.59	7.78
Services	16.01	-0.10	-0.29	9.57	6.53	1.99	6.92

*average estimates for food category do not include estimates for the basket of 22 food items.

As it can be seen from the table 4 above, aggregated estimates of the speed of convergence from the AR(1) model follow the same pattern as in panel unit root tests: services have the slowest speed of convergence and food commodities adjust faster to the average Ukrainian price level. The estimates of the speed of convergence from the TAR model are lower (higher by absolute value) than in AR(1) specification. As a result, half-lives from the TAR model are three times lower than in the AR(1) model. For all three subcategories of commodities average half-lives are in the range from 1 to 2 months. Half-lives from Levin and Lin test are consistently higher than the estimates from the TAR model.

Comparing the thresholds within the three categories of commodities, we can conclude that, as it was expected, food commodities have the lowest average threshold level of, approximately, 6%, while nonfood goods have the highest threshold of 8%. It might be surprising that services have a somewhat lower threshold level (7%) than nonfood category. However, it is worth mentioning here that the analysis of regional retail prices presented here considers only 4 services because of changed classification made by Derzhcomstat. These 4 services do not have the highest shares in household consumption and, therefore, is not a representative sample of services in Ukraine. In addition, the difference between the threshold levels for services and nonfood category is not big enough (only 1%) to give it some meaningful explanation.

As it is also can be seen from the table 4 above, the basket of 22 food goods has both the half-life and the threshold level lower than the whole category of food commodities. This can be explained by the fact that in addition to goods included in the food group, "dekadka" also contains such goods as carrot, apples, onion, and other goods that were not included into this analysis. Markets for these goods are very competitive and, therefore, drag down the estimates of half-life and threshold level for the whole basket.

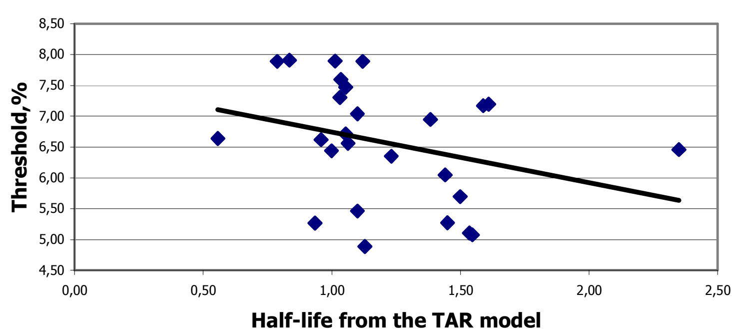
It might be also interesting to see what are differences not only between different commodities but also between different regions in Ukraine. How fast can retail prices in different oblasts of Ukraine adjust to the average Ukrainian level?

Table C5 (Appendix C) shows the average estimated coefficients from the TAR model for each location. Since calculation of a single estimate for each oblast of Ukraine is done to show in which oblasts of Ukraine arbitrage can be the most and the least profitable business, coefficients presented in table C5 (Appendix C) are the average coefficients for each location only across tradable goods (food and nonfood category, services are excluded).

Half-life for tradable goods from the TAR model for Ukrainian oblasts ranges from 0.56 months for Cherkaska oblast to 2.35 months for Odeska oblast. However, when oblasts are compared by the threshold level, Sumska oblast has the lowest average threshold level for tradable goods (4.89%) and Khersonska oblast has the highest average threshold (7.91%).

There is a clear negative relationship between the size of the threshold

Fig. 9. Relationship between the threshold levels and half-lives for Ukrainian regions (tradable goods)



and half-life of deviations of regional prices from the average Ukrainian level (fig. 9). This negative relationship is explained by the fact that when the threshold level is high, only very large deviations from the law of one price will appear in the subsample outside the band. These large deviations are more likely to damp out quickly. If the threshold level is small, then even small deviations of relative prices will be in the outside-the-band subsample. These small deviations will less likely be eliminated.

Taking into account transaction costs in the analysis of Ukrainian regional price differences changes results substantially compared with the panel unit root tests. Even though LLR test was not able to reject AR(1) model in favor of Band-TAR specification in some cases due to small univariate time series sample size, TAR model can still be considered as better than AR(1) model based on the significance of the coefficient λ^{out}, the speed of convergence for deviations from the law of one price outside the threshold band.

Results of the TAR model estimation confirm the conclusion of panel unit root tests that the speed of convergence is much faster for food commodities than for other categories. However, the magnitude of half-life of deviations from the equilibrium obtained from the TAR model are much lower than in panel unit root tests because the TAR model allows to distinguish between observations that fall inside the band of inaction and observations for which elimination of the shock effect is fast.

Thus, half-lives of deviations of relative price from equilibrium for food category range from 0.5 month for potato to 1.63 months for sour cream. Half-lives for nonfood and services are between 1 and 2 months. Thresholds vary from 3% to 15% for food, from 4% to 12% for nonfood, and from 5% to 9% for services.

Aggregation of the TAR estimation results for traded goods by location indicate that, on average, half-lives for Ukrainian regions vary from 0.5 months to 2 months with thresholds being in the range from 5% to 8%.

Chapter 7

CONCLUSIONS

The analysis of 30 regional retail prices for the period 1997:01-2002:12 in Ukraine relative to the average Ukrainian level shows that the law of one price holds when transaction costs are taken into account in the estimation procedure.

Even though, analysis of aggregated price indexes indicate that regional price differences can be persistent and may reach, approximately, 30% for some region pairs, panel unit root tests indicate that the fastest speed of adjustment can be expected for those tradable goods that have competitive markets and are not regulated.

Band-TAR specification uses more advanced analysis allowing for nonlinearity in the speed of convergence and bounds of inaction in which profits of arbitrage do not cover transaction costs. Partitioning sample into deviations inside and outside the bounds of inaction Band-TAR model estimation results substantially reduce half-lives of deviations of regional prices from the average Ukrainian level. Results of the TAR model suggest that regional prices do convergence to the average Ukrainian level with food goods having the fastest speed of convergence. Services category is shown to have three times slower speed of convergence compared to food category, which is line with finding of other studies.

Threshold levels are found to be in the range from 3% to 15% for different categories of goods and regions. Grouping of the TAR model results by location shows that those regions that have faster speed of convergence, on average, also have higher thresholds. Half-life of tradable goods from the

TAR model range from 0.56 months for Cherkaska oblast to 2.35 months for Odeska oblast.

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APPENDIX A. PANEL UNIT ROOT TESTS' ESTIMATION PROCEDURE.

A1. The Levin Lin Test (Levin and Lin, 1993)

Step 1. Subtract cross-section averages from the data.

$$\tilde{q}_{i,t} = q_{i,t} - \frac{\sum_{i=1}^{N} q_{i,t}}{N} \quad ,$$

where i is the subscript for cross-section, i $= 1, N$.

Step 2. Compute orthogonalized first differences and lagged levels for each cross-section, and normalize them by the estimated standard error.

2a. Regress $\Delta \tilde{q}_{i,t}$ on a constant and k_i lagged values of $\Delta \tilde{q}_{i,t}$, where k_i is a lag length. Let $\hat{e}_{i,t}$ denote the residuals from this regression.

2b. Regress $\tilde{q}_{i,t-1}$ on the same variables as in part 2a. Let $\hat{v}_{i,t-1}$ denote the residuals from this regression.

2c. Regress $\hat{e}_{i,t}$ on $\hat{v}_{i,t-1}$ without a constant. Let $\varepsilon_{i,t}$ denote the residuals from this regression. Normalize $\hat{e}_{i,t}$ and $\hat{v}_{i,t-1}$ by standard error of this regression $\hat{\sigma}_{ei}$. Denote the normalized values by $\tilde{e}_{i,t} = \frac{\hat{e}_{i,t}}{\hat{\sigma}_{e,i}}$ and $\tilde{v}_{i,t-1} = \frac{\hat{v}_{i,t-1}}{\hat{\sigma}_{ei}}$.

Step 3. Run the panel OLS regression $\tilde{e}_{i,t} = \beta \tilde{v}_{i,t-1} + u_{i,t}$

Step 4. The t-statistic of coefficient β has nonstandard distribution. However, Levin and Lin show that after some transformation such a t-statistic can be calculated that is asymptotically normally distributed under the null hypothesis and under the assumption that there is no contemporaneous correlation in errors.

A2. IPS test (Im, Pesaran, Shin, 1997)

Step 1. Substract cross-section averaged from the data (this step is identical to step 1 in Levin and Lin test).

Step 2. Run the augmented Dickey-Fuller regression for each cross-section: regress $\Delta\tilde{q}_{i,t}$ on a constant, $\tilde{q}_{i,t-1}$, and k_i lagged values of $\Delta\tilde{q}_{i,t}$, where k_i is a lag length. Let t_i denote the t-statistic of the coefficient of $\tilde{q}_{i,t-1}$ from the univariate ADF test.

Step 3. The IPS test statistic is:

$$\bar{t} = \frac{\sum_{i=1}^{N} t_i}{N}$$

Under the null hypothesis that each of the series contains a unit root and they are cross-sectionally independent, $\bar{t}$ a follow nonstandard distribution. Critical values are tabulated by Monte Carlo simulationsand reported in Im, Pesaran, and Shin (1997).

APPENDIX B. BAND-TAR MODEL ESTIMATION PROCEDURE (OBSTFELD AND TAYLOR, 1997)

H_1: Band-TAR model specification used in the analysis:

$$\Delta q_{i,rs,t} = \begin{cases} \lambda^{out}(q_{i,rs,t-1} - c) + e^{out}_t & \text{if } q_{i,rs,t-1} > c \\ \lambda^{in} q_{i,rs,t-1} + e^{in}_t & \text{if } c \leq q_{i,rs,t-1} \geq -c \quad \text{Band-TAR (1;2;1)} \\ \lambda^{out}(q_{i,rs,t-1} + c) + e^{out}_t & \text{if } -c > q_{i,rs,t-1} \end{cases}$$

where $e^{out}_t \sim N(0, \sigma^{out\,2})$, $e^{in}_t \sim N(0, \sigma^{in\,2})$, restriction $\lambda^{in}=0$ can be imposed.

H_0: AR1 model applied for the whole sample:

$$\Delta q_{i,rs,t} = \lambda_{i,rs} * q_{i,rs,t-1} + e_{i,rs,t} \qquad \text{(AR1)}^{1}$$

where $e_{i,rs,t} \sim N(0, \sigma^2)$.

The log likelihood ratio is:

$$LLR=2*(L_{TAR} - L_{AR1})$$

The objective is to maximize log likelihood ratio. A search algorithm:

Step1. Find the 25th and 75th percentiles[2] of series $|q_{i,rs,t}|$. Divide the interval from 75th to 25th percentiles into steps of 0.001 width. Let c_k denote corresponding thresholds (k=1,M). c_k are candidate thresholds. Put k=1.

---

[1] There is no constant in the specification because series $q_{i,rs,t}$ are already demeaned.

[2] Percentiles should be chosen on two grounds. First, on the basis of knowledge of what are the average transaction costs between regions in Ukraine. Second, taking into account the size of the time series sample for each cross-section some studies require that at least 30 percent of the sample were in any subsample. Obstfeld and Taylor (1997) have 180 observations and choose 10th and 90th percentiles as the bounds for a threshold in order to ensure that there are at least 18 observations in each part of the partition. Number of observations for each cross-section in the Ukrainian data used in this analysis is 72 for food and nonfood commodities and there are 60 observations for each cross-section of services. Consequently, 25th and 72th percentiles (at least 18 observations in each subsample) are chosen as bounds for thresholds when food and nonfood goods are analyzed. Corresponding bounds for services are the 70th and 30th percentiles. In order to detect whether the estimates of TAR model are sensitive to the choice of bounds, coefficients were estimated for 20/80, 25/75, and 30/70 percentiles as bounds for the threshold parameter. λ^{out} has not changed significantly in all three cases. Coefficient λ^{in} was significantly

Step 2. Choose a threshold c to be equal to the highest value of c_k. At the first step this will be the value of the 75th percentile.

Step 3. Divide the sample into two subsamples: with observations inside and outside the band. Transform the values $q_{i,rs,t-1}$ into ($q_{i,rs,t-1}$ - c) if $q_{i,rs,t-1}>c$ and into ($q_{i,rs,t-1}$ + c) if $-c> q_{i,rs,t-1}$. Observations that are inside the band are not transformed.

Step 4. Calculate L_{TAR} by estimating two regressions on partitioned samples with observations inside and outside the band by maximum likelihood or by OLS.

$$L_{TAR}= L_{outside} + L_{inside}$$

Then calculate L_{AR} by estimating AR(1) model for the whole sample.

$$LLR_k = 2*(L_{TAR} - L_{AR1})$$

Step 5. Increase k by one and set $c_k=(c_{previous} - 0.001)$.

Step 6. Repeat steps 3-5 until c_k reaches the lower bound, the value of the 25th percentile. Find the threshold c_k at which LLR is maximized (example is shown in figure 4 in the text). Find the corresponding values of λ^{out} and λ^{in}.

Since the threshold parameter c is not identified under H_0, LLR statistic follows nonstandard distribution. **Monte Carlo simulation** should be done to find critical values of this distribution.

Step 1. Estimate the AR(1) model on the actual data (q_1, q_2, …., q_T).

Step 2. Generate 600 simulations of this model. Start each with $q_{-50}=0$, end with q_T. Discard the first 50 observations to avoid bias in initial values.

Step 3. For each simulation, estimate the TAR model and find maximized LLR_i . Order the obtained 600 values of LLR_i and compare your real calculated LLR statistic found in step 6 with simulated values. Find p-value of LLR.

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different in some cases but it was statistically insignificant. Taking into account that λ^{in} is not important for the analysis of the results of the TAR model and in some studies is set to be equal to 0, we may conclude that the choice of threshold bounds does not influence final results.

APPENDIX C. PANEL UNIT ROOTS AND TAR MODEL

Table C1. The list of commodities used in the analysis of nominal retail prices.

Good	Period	Description	Good	Period	Description
Food			**Nonfood**		
Sour cream	1997:01-2002:12	1 kg	Boots for men	1997:01-2002:12	1 pair
Flour (wheat)	1997:01-2002:12	1 kg	Boots for women	1997:01-2002:12	1 pair
Milk	1997:01-2002:12	1 litre	Refrigerator	1997:01-2002:12	1 item
Sugar	1997:01-2002:12	1 kg	Skirt for women	1997:01-2002:12	1 item
Beef	1997:01-2002:12	1 kg	Soap (household)	1997:01-2002:12	200 gram
Butter	1997:01-2002:12	1 kg	Sofa (bed)	1997:01-2002:12	1 item
Oil (sunflower)	1997:01-2002:12	1 kg	Stockings for women	1997:01-2002:12	1 pair
Potato	1997:01-2002:12	1 kg	Stockings for children	1997:01-2002:12	1 pair
Poultry	1997:01-2002:12	1 kg	Jacket for children	1997:01-2002:12	1 item
Rice	1997:01-2002:12	1 kg	**Services**		
Beet	1997:01-2002:12	1 kg	Shoes repair	1997:01-2001:12	
Cabbage	1997:01-2002:12	1 kg	Photo services	1997:01-2001:12	6 photos for documents
Semolina	1997:01-2002:12	1 kg	Laundry and ironing	1997:01-2001:12	price for 1 kg of linens
Cheese (hard)	1997:01-2002:12	1 kg	Haircut for men	1997:01-2001:12	
Horilka (alcohol)	1997:01-2002:12	1 litre			
Fat (salo)	1997:01-2002:12	1 kg			
Mineral water	1997:01-2002:12	1 litre			
Basket of 22 food goods	1995:01-1998:12	total cost			

Table C2. Summary of panel unit root tests' results.

	Levin and Lin test		IPS test		Levin and Lin test		IPS test
Good	**Coeffcient λ**[1]	**Half-life**[1]	**p-value**[2]	**Good**	**Coeffcient λ**[1]	**Half-life**[1]	**p-value**[2]
Food (#of observations =1534 for each panel)				**Nonfood** (#of observations=1534 for each panel)			
Sour cream	**-0.09316** (0.9673)[3]	**7.09**	(0.255)	Boots for men	**-0.10832** (0.9943)	**6.05**	(0.116)
Flour	**-0.15720** (1.0000)	**4.05**	(0.007)	Boots for women	**-0.08829** (0.9214)	**7.5**	(0.265)
Milk	**-0.18459** (1.0000)	**3.40**	(0.451)	Refrigerator	**-0.10993** (0.9689)	**5.95**	(0.250)
Sugar	**-0.23308** (1.0000)	**2.61**	(0.001)	Skirt for women	**-0.08862** (0.9998)	**7.47**	(0.183)
Beef	**-0.09382** (1.0000)	**7.04**	(0.475)	Soap (household)	**-0.10655** (0.9921)	**6.15**	(0.160)
Butter	**-0.17985** (1.0000)	**3.50**	(0.091)	Sofa (bed)	**-0.08644** (1.0000)	**7.67**	(0.171)
Oil	**-0.16229** (0.8246)	**3.91**	(0.003)	Stockings for women	**-0.08860** (0.9996)	**7.47**	(0.067)
Potato	**-0.34702** (1.0000)	**1.63**	(0.000)	Stockings for children	**-0.08765** (0.8758)	**7.56**	(0.264)
Poultry	**-0.18386** (1.0000)	**3.41**	(0.015)	Jacket for children	**-0.09611** (0.9502)	**6.86**	(0.356)
Rice	**-0.18079** (1.0000)	**3.48**	(0.000)	**Average**	**-0.095612 (0.000)**[5]	**6.9**	
Beet	**-0.31497** (1.0000)	**1.83**	(0.000)	**Services** (#of observations =1222 for each panel)			
Cabbage	**-0.45100** (1.0000)	**1.16**	(0.000)				
Semolina	**-0.16380** (0.9999)	**3.87**	(0.002)	Shoes repair	**-0.07378** (0.9762)	**9.04**	(0.260)
Cheese	**-0.19630** (1.0000)	**3.17**	(0.015)	Photo services	**-0.07407** (0.0294)	**9.01**	(0.000)
Horilka	**-0.13385** (0.9297)	**4.82**	(0.091)	Laundry and ironing	**-0.05503** (0.7298)	**12.25**	(0.357)
Fat	**-0.14722** (1.0000)	**4.35**	(0.132)	Haircut for men	**-0.07667** (0.8422)	**8.69**	(0.152)
Mineral water	**-0.06531** (0.5748)	**10.26**	(0.158)	**Average**	**-0.069887 (0.000)**[5]	**9.57**	
Basket of 22 food goods	**-0.30110** (1.0000)	**1.93**	(0.001)				
Average[4]	**-0.193418 (0.000)**[5]	**3.22**					

[1] Speed of convergence and half-life for Levin and Lin tests are shown for the case when 12 lags are included.
[2] P-value of IPS statistic is shown for the case when 12 lags are included.
[3] P-values of coefficients are given in brackets.
[4] Average estimate for food do not include coefficient of the basket of 22 food goods.
[5] Average coefficient is very significant because the variance of such coefficient is very low.

Table C3. Full results of panel unit root tests.

	Levin Lin test					IPS test
Good	**4 lags** ***Coeffcient λ***	**6 lags** ***Coeffcient λ***	**8 lags** ***Coeffcient λ***	**12 lags** ***Coeffcient λ***	**Half-life (for12 lags)**	**p-value (12 lags)**
FOOD						
# of observations	**1742**	**1690**	**1638**	**1534**		
Sour cream	-0.10331 [.01100907][1] (0.0092)[2]	-0.09536 [.01147111] (0.2978)	-0.09570 [.01222971] (0.6998)	**-0.09316** [.01288916] (0.9673)	7.09	(0.255)
Flour	-0.12644 [.01375805] (0.1666)	-0.12881 [.0147949] (0.7079)	-0.14307 [.0161395] (0.8794)	**-0.15720** [.02002601] (1.0000)	4.05	(0.007)
Milk	-0.19795 [.02040345] (0.0952)	-0.16362 [.02248481] (0.9983)	-0.14793 [.02362927] (1.0000)	**-0.18459** [.02712373] (1.0000)	3.40	(0.451)
Sugar	-0.33570 [.0226813] (0.0000)	-0.29012 [.02520816] (0.9617)	-0.26732 [.0259207] (1.0000)	**-0.23308** [.0286023] (1.0000)	2.61	(0.001)
Beef	-0.10893 [.01272125] (0.3027)	-0.10541 [.0136497] (0.8849)	-0.09237 [.01437868] (0.9995)	**-0.09382** [.01612231] (1.0000)	7.04	(0.475)
Butter	-0.20868 [.01736885] (0.0026)	-0.18976 [.01881682] (0.7783)	-0.16452 [.01995558] (0.9999)	**-0.17985** [.02243385] (1.0000)	3.50	(0.091)
Oil	-0.16555 [.01463974] (0.0006)	-0.15587 [.01530199] (0.1808)	-0.13970 [.01490212] (0.7364)	**-0.16229** [.01577087] (0.8246)	3.91	(0.003)
Potato	-0.36669 [.02620736] (0.0044)	-0.37202 [.0287955] (0.7866)	-0.33515 [.03090259] (1.0000)	**-0.34702** [.034313] (1.0000)	1.63	(0.000)
Poultry	-0.14175 [.01408163] (0.2875)	-0.13829 [.01509496] (0.9462)	-0.15159 [.01604927] (0.9737)	**-0.18386** [.01938436] (.0000)	3.41	(0.015)
Rice	-0.17123 [.01432736] (0.0004)	-0.18047 [.01574627] (0.0579)	-0.18858 [.01722008] (0.5693)	**-0.18079** [.02046362] (1.0000)	3.48	(0.000)
Beet	-0.32439 [.02253343] (0.0000)	-0.33249 [.02506342] (0.2140)	-0.30225 [.02766297] (0.9998)	**-0.31497** [.03125774] (1.0000)	1.83	(0.000)
Cabbage	-0.42546 [.02622422] (0.0000)	-0.41295 [.02962342] (0.7373)	-0.41188 [.03308564] (1.0000)	**-0.45100** [.03876639] (1.0000)	1.16	(0.000)
Semolina	-0.14629 [.01348473] (0.0063)	-0.14889 [.01460879] (0.2627)	-0.15643 [.01594065] (0.7497)	**-0.16380** [.01892426] (0.9999)	3.87	(0.002)
Cheese	-0.20552 [.01535294] (0.0000)	-0.17401 [.01674156] (0.7323)	-0.15638 [.01817491] (1.0000)	**-0.19630** [.02216346] (1.0000)	3.17	(0.015)
Horilka	-0.11326 [.01302199] (0.1181)	-0.10028 [.01372602] (0.8766)	-0.10609 [0.9285] (.0145385)	**-0.13385** [.0165268] (0.9297)	4.82	(0.091)
Fat	-0.14592 [.01504903]	-0.14002 [.01665236]	-0.13253 [.01757909]	**-0.14722** [.02013496]	4.35	(0.132)

	(0.6991)	(0.9997)	(1.0000)	(1.0000)		
Mineral water	-0.06619 [.00814281] (0.0024)	-0.05523 [.00813654] (0.1475)	-0.05798 [.00850288] (0.1901)	**-0.06531** [.00966481] (0.5748)	10.26	(0.158)
Basket of 22 food goods	-0.30477 [.02789535] (0.3929)	-0.29024 [.02956037] (0.9660)	-0.30270 [.03045593] (0.9988)	**-0.30110** [.03360353] (1.0000)	1.93	(0.001)
Average[3]				**-0.193418** **[0.00564]** **(0.000)**[3]	**3.22**	
	Levin Lin test					**IPS test**
Good	**4 lags**	**6 lags**	**8 lags**	**12 lags**	**Half-life**	**p-value**
NONFOOD						
# of observations	**1742**	**1690**	**1638**	**1534**		
Boots for men	-0.09775 [.01082476] (0.0518)	-0.09636 [.01154348] (0.4467)	-0.10106 [.01236375] (0.7283)	**-0.10832** [.0141638] (0.9943)	6.05	(0.116)
Boots for women	-0.07659 [.00970729] (0.1004)	-0.07199 [.01030312] (0.6706)	-0.06953 [.01082109] (0.9362)	**-0.08829** [.01218404] (0.9214)	7.5	(0.265)
Refrigerator	-0.08496 [.01129623] (0.2642)	-0.08946 [.01207866] (0.5126)	-0.08921 [.01285823] (0.8511)	**-0.10993** [.01509511] (0.9689)	5.95	(0.250)
Skirt for women	-0.07439 [.00960157] (0.5944)	-0.07447 [.01009183] (0.8821)	-0.07961 [.01067453] (0.9495)	**-0.08862** [.01235874] (0.9998)	7.47	(0.183)
Soap (household)	-0.07801 [.0102308] (0.4341)	-0.08072 [.01101699] (0.7985)	-0.08953 [.01176631] (0.8320)	**-0.10655** [.01408804] (0.9921)	6.15	(0.160)
Sofa (bed)	-0.07758 [.01134244] (0.7898)	-0.07406 [.0120173] (0.9853)	-0.07619 [.01280053] (0.9977)	**-0.08644** [.01479895] (1.0000)	7.67	(0.171)
Stockings for women	-0.07901 [.00983041] (0.3276)	-0.08400 [.01065277] (0.6792)	-0.09473 [.01154863] (0.7606)	**-0.08860** [.01280163] (0.9996)	7.47	(0.067)
Stockings for children	-0.06087 [.00909529] (0.5644)	-0.06468 [.00967975] (0.7266)	-0.06901 [.01028188] (0.8144)	**-0.08765** [.01191684] (0.8758)	7.56	(0.264)
Jacket for children	-0.07991 [.01008755] (0.2098)	-0.08245 [.0106608] (0.4475)	-0.08440 [.01130604] (0.7127)	**-0.09611** [.01295593] (0.9502)	6.86	(0.356)
Average				**-0.095612** **[0.00447]** **(0.000)**[3]	**6.9**	

	Levin Lin test					IPS test
Good	**4 lags**	**6 lags**	**8 lags**	**12 lags**	**Half-life**	**p-value**
SERVICES						
# of observations	**1430**	**1378**	**1326**	**1222**		
Shoes repair	-0.05189 [.00901837] (0.6732)	-0.05906 [.00970056] (0.6249)	-0.06109 [.01045138] (0.8744)	**-0.07378** [.01232824] (0.9762)	9.04	(0.260)
Photo services	-0.05283 [.00693971] (0.0014)	-0.06348 [.00749265] (0.0002)	-0.06987 [.00788939] (0.0001)	**-0.07407** [.00881447] (0.0294)	9.01	(0.000)
Laundering and ironing	-0.04925 [.0076607] (0.0580)	-0.04825 [.00788888] (0.1757)	-0.05005 [.00824712] (0.2702)	**-0.05503** [.00942332] (0.7298)	12.25	(0.357)
Haircut for men	-0.05636 [.0085494] (0.3583)	-0.06232 [.00919714] (0.3659)	-0.07171 [.00979895] (0.3077)	**-0.07667** [.01109504] (0.8422)	8.69	(0.152)
A verage				**-0.069887 [0.005253] (0.000)[3]**	**9.57**	

[1] Standard deviations of coefficients

[2] P-values of coefficients

[3] Average coefficient is very significant because the variance of such coefficient is very low.

Table C4. TAR model results (grouped by commodity type).

#	Good	T (total)	T (outside)	LLR	Lambda AR(1)	Lambda TAR	Half-life (AR1)	Half-life (TAR)	Threshold, %	Rejection, %
1	Sour cream	72	31	7.59	-0.13	-0.35	4.83	1.63	6.62	19
2	Flour	72	34	13.71	-0.20	-0.51	3.18	0.98	3.64	54
3	Milk	72	35	31.93	-0.28	-0.47	2.14	1.08	7.17	38
4	Sugar	72	33	12.38	-0.34	-0.65	1.66	0.65	3.10	38
5	Beef	72	36	12.00	-0.16	-0.39	3.93	1.42	6.20	46
6	Butter	72	37	7.82	-0.24	-0.49	2.48	1.04	2.93	15
7	Oil	72	35	11.76	-0.20	-0.40	3.07	1.35	5.21	58
8	Potato	72	40	15.10	-0.45	-0.74	1.17	0.51	6.70	54
9	Poultry	72	36	8.74	-0.20	-0.38	3.08	1.44	5.01	27
10	Rice	72	35	12.27	-0.21	-0.40	2.97	1.35	3.11	42
11	Beet	72	32	23.60	-0.35	-0.66	1.59	0.64	12.54	77
12	Cabbage	72	32	9.74	-0.40	-0.92	1.37	0.28	14.59	31
13	Semolina	72	32	12.86	-0.20	-0.42	3.09	1.27	3.29	58
14	Cheese	72	33	9.04	-0.24	-0.50	2.49	0.99	3.90	27
15	Horilka	72	31	18.15	-0.21	-0.59	2.95	0.78	5.29	77
16	Fat	72	34	8.63	-0.19	-0.41	3.26	1.32	5.36	23
17	Mineral water	72	31	16.58	-0.14	-0.36	4.76	1.56	6.96	46
	Basket of 22 food goods	**48**	**23**	**5.94**	**-0.38**	**-0.73**	**1.47**	**0.53**	**2.56**	**23**
	Food		**34**	**13.64**	**-0.24**	**-0.51**	**2.48**	**0.98**	**5.93**	**43**
18	Boots for men	72	31	18.40	-0.14	-0.37	4.64	1.51	8.64	54
19	Boots for women	72	35	16.35	-0.12	-0.34	5.51	1.67	9.91	65
21	Refrigerator	72	32	15.10	-0.17	-0.41	3.67	1.30	6.40	50
22	Skirt for women	72	32	13.44	-0.13	-0.30	5.18	1.94	8.18	54
23	Soap (household)	72	33	13.19	-0.13	-0.33	4.90	1.72	5.20	38
24	Sofa (bed)	72	33	22.23	-0.15	-0.47	4.36	1.09	8.07	65
25	Stockings for women	72	35	19.17	-0.13	-0.27	4.98	2.19	7.45	50
26	Stockings for children	72	36	21.13	-0.11	-0.30	6.08	1.93	4.42	58
27	Jacket for children	72	32	17.21	-0.13	-0.38	4.92	1.46	11.96	58
	Nonfood		**33**	**17.36**	**-0.13**	**-0.35**	**4.83**	**1.59**	**7.78**	**55**
27	Shoes repair services	60	28	17.78	-0.10	-0.29	6.54	2.02	5.09	58
28	Photo services	60	27	21.40	-0.11	-0.33	6.13	1.72	6.89	65
29	Laundry and ironing	60	31	14.13	-0.09	-0.28	6.95	2.10	8.69	42
30	Haircut for men	60	28	10.73	-0.10	-0.27	6.56	2.16	7.05	31
	Services		**29**	**16.01**	**-0.10**	**-0.29**	**6.53**	**1.99**	**6.92**	**49**

*Average estimate for food category does not include the estimate for the basket of 22 food commodities.

** Coefficient λ from TAR model is λ^{out}. λ^{in} are not reported because they are insignificant.

***Coefficients λ from AR(1) and TAR models are average coefficients for 26 cross-sections for each commodity. Their p-values are not reported because they are all very significant (because variances of average coefficients are very low).

****The column "Rejection,%" shows in how many cross sections AR(1) model was rejected in favor of Band-TAR model by LLR test

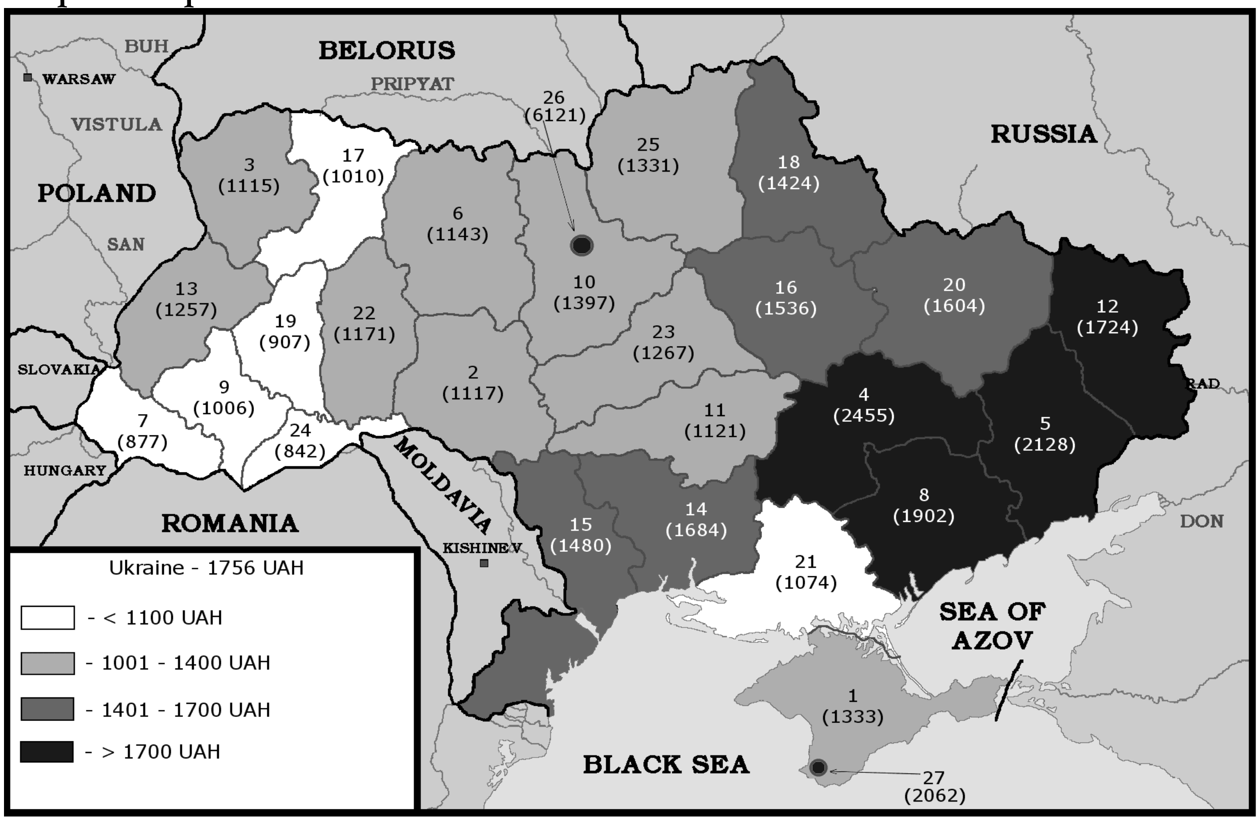
Table C5. TAR model results for traded goods (grouped by location).

#	Oblast	T (total)	T (outside)	LLR	Lambda AR(1)	Lambda TAR	Half-life (AR1)	Half-life (TAR)	Threshold, %
1	Crimea	72	33	10.57	-0.15	-0.38	4.25	1.44	6.05
2	Vinnytska	72	31	16.44	-0.18	-0.50	3.48	1.00	6.44
3	Volynska	72	31	15.96	-0.20	-0.49	3.03	1.04	7.60
4	Dniproperovska	72	33	11.31	-0.15	-0.47	4.29	1.10	7.04
5	Donetska	72	38	11.49	-0.17	-0.36	3.71	1.55	5.08
6	Zhytomyrska	72	29	20.05	-0.15	-0.35	4.40	1.59	7.17
7	Zakarpatska	72	34	10.03	-0.18	-0.39	3.52	1.38	6.95
8	Zaporizka	72	36	15.94	-0.18	-0.37	3.47	1.50	5.69
9	Ivano-Frankivska	72	34	11.68	-0.15	-0.35	4.19	1.61	7.20
10	Kyivska	72	36	12.97	-0.18	-0.36	3.49	1.54	5.11
11	Kirovogradska	72	39	9.64	-0.26	-0.48	2.35	1.05	6.71
12	Luganska	72	31	13.74	-0.24	-0.52	2.55	0.93	5.27
13	Lvivska	72	29	16.32	-0.16	-0.50	3.92	1.01	7.89
14	Mykolaivska	72	35	8.93	-0.22	-0.48	2.80	1.06	6.56
15	Odeska	72	39	12.93	-0.13	-0.26	4.92	2.35	6.46
16	Poltavska	72	34	16.54	-0.23	-0.47	2.66	1.10	5.46
17	Rivnenska	72	35	16.49	-0.18	-0.48	3.43	1.05	7.47
18	Sumska	72	38	12.55	-0.23	-0.46	2.60	1.13	4.89
19	Ternopilska	72	32	13.21	-0.24	-0.46	2.51	1.12	7.89
20	Kharkivska	72	29	24.01	-0.18	-0.49	3.50	1.03	7.31
21	Khersonska	72	35	16.06	-0.29	-0.56	2.05	0.84	7.91
22	Khmelnytska	72	34	24.61	-0.22	-0.43	2.85	1.23	6.35
23	Cherkaska	72	32	19.55	-0.36	-0.71	1.53	0.56	6.64
24	Chernivetska	72	34	17.79	-0.25	-0.51	2.38	0.96	6.62
25	Chernigivska	72	32	16.03	-0.29	-0.59	1.99	0.79	7.89
26	Kyiv	72	33	13.33	-0.17	-0.38	3.84	1.45	5.27

* Coefficient λ from TAR model is λ^{out}. λ^{in} are not reported because they are insignificant.

Coefficients λ from AR(1) and TAR models are average coefficients for 26 tradable (food and nonfood, services are excluded**) goods for each location. Their p-values are not reported because they are all very significant (because variances of averaged coefficients are very low).

Graph 1. Map of Ukraine.



1–Crimea autonomy; oblasts: 2–Vinnytsya, 3–Volyn, 4–Dnipropetrovsk, 5–Donetsk, 6–Zhytomyr, 7–Transkarpatian, 8–Zaporizzya, 9–Ivano-Frankivsk, 10–Kyiv, 11–Kirovograd, 12–Lugansk, 13–Lviv, 14–Mykolaiv, 15–Odesa, 16–Poltava, 17–Rivne, 18–Sumy, 19–Ternopil, 20–Kharkiv, 21–Kherson, 22–Khmelnytskiy, 23–Cherkasy, 24–Chernivtsi, 25–Chernigiv; 26–Kyiv city, 27–Sevastopol city.