Integration of the US goods market, 2001–2015

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Abstract

A spatially dispersed market for a tradable good is deemed integrated if there are no barriers to trade between its spatial segments (except for geographical barriers, namely distances between the segments). However, perfectly integrated markets are not a common case; real markets deviate to some extent from this ideal state. Therefore, estimating a degree of integration is more helpful then an answer of the type “all or nothing” (whether the market is integrated or not integrated). In an integrated market, price for a good is determined in the national market as a whole, not depending on demand in its spatial segments. Hence, a dependence of local price on local demand (controlling for transportation costs) indicates a deviation from perfect integration, and its “strength” can measure the degree of market integration. Based on this idea, we estimate the annual integration degrees of the US market for an aggregated good (grocery basket) over 15 years, 2001–2015. The spatial segments are cities; our sample covers 66 cities from 39 states of the US. The results suggest that the US market is not perfectly integrated; however, the integration degree of the US market is fairly stable over time. We also compare results for the US with results of a similar study for Russia. With a reservation that the empirical material is not fully comparable, we can conclude that the US market is integrated more strongly than the Russian market and that the integration degree in Russia is more volatile.

Keywords Spatial market integration · Price dispersion · Law of one price · United States

JEL Classification L81 · R15 · R19

1 Introduction

An intuitive definition of integrated market for goods was proposed almost two centuries ago, in 1838, as "an entire territory of which the parts are so united by the relations of unrestricted commerce that prices take the same level throughout with ease and rapidity" (Cournot, 1971, pp. 51–52). In contemporary terms, this means that the law of one price holds in the integrated market, deviations from it being quickly eliminated. Unfortunately, real markets fall far short of this ideal pattern.

First of all, the parts of a territory under consideration (hereafter, locations) are spatially separated, so causing transportation costs that are included into commodity prices. A weak version of the law of one price allows prices to differ between locations by no more than transportation costs. However, controlling for these costs does not fully solve the problem. The point is that prices include non-tradable components due to wages and rents that can differ across locations because of poor (if any) integration of related markets – labor market and real estate market. One more reason is that opportunities for spatial arbitrage are not fully exploited because of incomplete information, long-term contracts between members of the sales chains, etc. In total, all these and other market frictions can be characterized as the arbitrage transaction costs per unit of good. When the difference in prices between two locations is less than the arbitrage transaction costs, arbitrage is unprofitable and does not occur.

Thus, any real goods market is segmented to a greater or lesser extent. Therefore, an assessment of a market as integrated or not integrated is insufficient. Instead of such a binary measure, it is more adequate to use a continuous measure indicating the degree of integration (or, by the same token, segmentation: the lesser the market segmentation, the stronger the integration).[1] In doing so, it is reasonable to consider arbitrage transaction costs less transportation costs. This implies that the spatial separation of locations is not deemed an impediment to integration. In other words, a benchmark of perfect integration (zero segmentation) is the fact that differences in prices between locations do not exceed transportation costs.

Certainly, it is not possible to estimate arbitrage transaction costs because they are not fully observable (especially their components caused by institutional reasons). Some indirect way of measuring the degree of market integration should be used instead. In this paper, we apply the way proposed by Gluschenko (2004). It is based on the following idea. In the integrated market, price for

[1] As Fackler and Goodwin (2001, p. 978) note, "In general, market integration refers to a measure of degree rather than a specific relationship. At one extreme are completely separated markets and at the other are perfectly integrated markets".

a tradable good in a location is determined in the national market as a whole, not depending on demand in this location. Hence, a dependence of local price on local demand – controlling for transportation costs – indicates a deviation from the law of one price in its weak version; the "strength" of the dependence can measure the degree of market segmentation. In turn, local demand depends on local per capita income, which makes it possible to dispense with unavailable data on demand. The elasticity of price difference (net transportation costs) with respect to income difference measures the degree of market segmentation.

Exploiting this methodology, we consider a market for an aggregated good (grocery basket) in the US. Locations are represented by 68 cities from 39 states of the US. The variable to be analyzed is the price differential, the logarithm of the ratio of the grocery basket costs of in two cities. The independent variables are the income differential and logarithm of distance between these cities (the latter being a proxy of transportation costs). Thus, the coefficient on the income differential is the measure of the segmentation degree in a market consisting of two cities. Our regression estimates the degree of segmentation over all pairs of cities for a given year. Running such regression for every year from 2001 to 2015, we obtain a dynamic pattern of market integration over 15 years,.

The first attempt of estimating US market integration with the use of the above methodology was made by Gluschenko and Karchevskaya (2010). In fact, it was a test for operability of the model. Their analysis relates to one period only (the first quarter of 2000) with a small spatial sample covering 29 cities. In this study, we obtain a much richer pattern, both in space and time.

The results obtained suggest, as might be expected, that the US market is not perfectly integrated. The degree of market integration is fairly stable over time, with no trend observed. However, there is no a standard by which to judge the strength of integration by the segmentation degrees obtained (other than zero segmentation indicating perfect integration). To solve this problem, we compare results for the US with results of a similar study for Russia. With a reservation that the empirical material is not fully comparable, we can conclude that the US market is integrated more strongly than the Russian market and that the integration degree in Russia is more volatile.

The rest of the paper is organized as follows. Section 2 reviews empirical studies of the US market integration. Section 3 presents methodology of the analysis. Section 4 describes data employed. Section 5 reports results of the analysis and compares them with estimates obtained for the Russian market. Section 6 summarizes the study.

2 Studies of market integration in the US

General publications on market integration, Fackler and Goodwin (2001) and Barrett (2008), note that there is no universally accepted definition of the term. The most widespread approach is just characterization of integration in the binary manner mentioned in Introduction, based on testing for the law of one price: a market is deemed integrated if the law holds and not integrated if it does not. It should be noted, however, that this is not clearly summarized in some cases; the authors only state whether the law holds or not.

Parsley and Wei (1996) apply this approach, analyzing time series of prices for 51 commodities (41 goods among them) over 1975–1992 with a quarterly frequency across 48 US cities. The source of their data is the ACCRA cost of living index.[2] The analysis lies in testing the law of one price with the use of standard AR(1) model and a panel unit root test. Each panel contains time series of price for an individual commodity in all cities. Thus, the analysis answers the question whether a given market is or is not integrated. Most markets for tradable goods are found to be integrated. At the same time, 5 out of 41 of such markets turn out to be not integrated; in a model with city and time fixed effects, this figure increases to 24. However, a reservation is needed here. A feature of the analysis is that the prices are presented as relative to those in the benchmark city, New Orleans. As a consequence, the pattern obtained tells about integration of New Orleans with the rest of cities, but is silent as to the integration of these cities with each other. Chmelarova and Nath (2010), using annual consumer price index (CPI) data for 17 US cities over 1918–2007, find the results of testing the law of one price to depend crucially on the choice of benchmark city.

O'Connell and Wei (2002) use substantially the same data as Parsley and Wei (1996), however, with smaller numbers of commodities (48) and cities (24); the time span under consideration is the same. The absolute prices in commodity panels are converted to relative prices with the use of the average price across all cities (rather than a price in some city) as the numeraire. One more difference with Parsley and Wei's study is more sophisticated analysis. It involves a GLS panel unit root test for the AR(1) model and estimating a threshold autoregression (TAR) model as well as an exponential smooth threshold autoregression (ESTAR) model. Among 35 markets for goods that are not classified as locally-produced, the law of one price does not hold in only two. Along with this, it is found that for approximately 60% of such goods sustained price gaps persist, evidencing that the law of one price holds in its weak version. Estimates of the threshold models reveal nonlinearities in the behavior of prices.

[2] This source of data on prices in the US is discussed in the next section.

Yazgan and Yilmazkuday (2011) analyze prices for 48 commodities across 52 US cities drawn from the ACCRA cost of living index; the time span is 1990–2007 with a quarterly frequency. In contrast to the above studies that use an arbitrary benchmark for constructing relative prices, relative prices are constructed for all 1326 pairs of cities, following Pesaran (2007). The analysis exploits four unit root tests applied to test stationarity of every individual relative price. Thus, it reveals whether a submarket consisting of given pair of cities is integrated. Choosing results of testing that give the maximum figure, the portion of integrated pairs is 87.2% for non-perishable goods and 83.3% for perishable goods.

An influential article by Cechetti et al. (2002) was one of the first to use long-span price series. Their data panel includes CPI over 1918–1995 for 19 US cities with the cross-sectional mean as a numeraire. The law of one price is found to hold in the whole time span. However, both applied unit root tests reject this in subperiods 1918–1955 and 1936–1955. This study gave rise to a series of other ones that revisited its analysis. Nath and Sarkar (2009) reexamine the study by correcting for both aggregation and small-sample biases and obtain a better pattern, namely, faster convergence to the law of one price. Basher and Carrion-i-Silvestre (2011) take account of structural breaks in the CPI series. With this innovation, convergence turns out even more fast. Sonora (2008) performs an alternative analysis of the same data, testing individual series (relative to Chicago) rather than the CPI panel. The number of cities where the law of one price holds is considerable, although not all applied tests find evidence in favor of the law in all series. A caveat similar to that made above with respect to Parsley and Wei's (1996) results is needed, though, about the crucial role of the choice of the benchmark city.

Harrison (2023) explores spatial market integration in the US over a very long period of time, 200 years (1750–1949). Because of different availability of historical price data, he breaks the analysis into five distinct subperiods, each with its own set of goods and locations. Price time series are decomposed (with the use of state space method) into three components: a trend, to measure convergence; a cycle, to measure efficiency; and a seasonal, to measure intertemporal smoothing. Convergence of prices between locations are deemed improvements in market integration. The convergence results evidence that prices converged prior to the American Revolution (started in 1775), did not converge after the adoption of the Constitution of the US (in 1789), and stagnated rather than diverged during the early 1800s. Then convergence occurred again, mostly accomplished by 1840. Constant convergence is found during 1855–1909, which suggests a large role played by railroads in improving arbitrage. In the final period (1910–1949), convergence occurred across all regions, albeit with temporary divergences caused by the nominal distortions in government-set rail

rates during the inflation after World War I and the Great Depression deflation.

A number of studies deal with markets for individual goods, for example, the gasoline market (Paul et al. 2001; Holmes et al. 2013; Barassi et al. 2025) and the market for such a specific good as cannabis (Goodwin 2025). As regards the gasoline market, although the listed studies differ in time spans (1983–1998, 1989–2011, and 2003–2019), types of locations (US Petroleum Administration Defense Districts – PADD, states and cities), and econometric techniques, their results evidence in favor of the law of one price in most cases. However, Paul et al. (2001) note that perfect market integration is rejected in all but a few cases. Goodwin's (2025) data covers 15 US states (a group of states that have legalized cannabis use) and the period of 2015–2021 with a weekly frequency. Using semiparametric models of spatial price linkages among US states, he finds that California cannabis markets (leading cannabis exporters) are largely integrated with states across the nation.

It is worth mentioning an important "sideline" of studies that do not directly address market integration but examine a specific impediment to it, the border effect. Its impact on prices has been discovered by Engel and Rogers (1996).[3] They apply a cross-sectional approach to examine price volatility of 14 goods (from disaggregated CPI) among pairs of 23 cities of the US and Canada. The finding is that the variation of prices of similar goods is much higher for two cities located in different countries than for two equidistant cities in the same country. The border effect is equivalent to a huge increase in the distance (compared to the actual distance) between cities on different sides of the US-Canada border. The seminal study due to Engel and Rogers (1996) has found a number of followers. Parsley and Wei (2001) examine the US-Japan border effect. As can be expected, it turns out much greater than the effect of US-Canada border. However, Gorodnichenko and Tesar (2009) show that the US-Canada border effect is entirely driven by intranational differences in the distribution of prices within the US and Canada. To shed light on this issue, Yilmazkuday (2012) estimates the effects of state borders within the US. The results suggest that these borders have positive and significant effects on price dispersion across the US cities; the distance equivalent of the state border effect is, on average, about 3,344 miles.

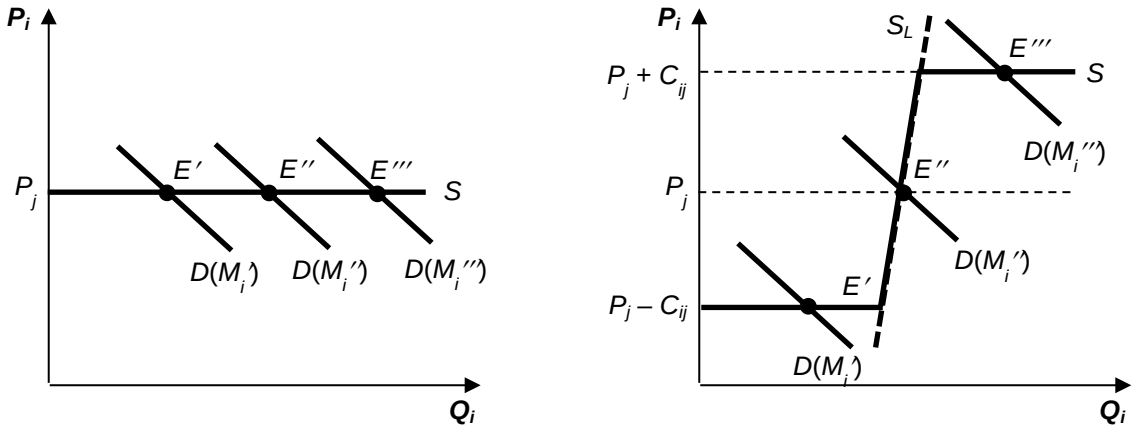
This review is, of course, not exhaustive. However, other studies not included in the review provide qualitatively similar patterns. In general, the results show that the US goods market is fairly strongly integrated, but at the same time it cannot be considered perfectly integrated.

[3] Their article was preceded by McCallum (1995), who examined the effect of border on shipments of goods between and within the US and Canada.

3 Methodology

Let a market for a tradable good consists of a set of spatial segments $\{k\}$; p_k stands for the price of the good in segment k; m_k is personal income per capita; q_k is the quantity of the good. Capital letters denote logarithms of corresponding variables, e.g., $P_k = \log(p_k)$. Consider a two-segment market $\{k\} = \{i, j\}$. Segment j can be interpreted as the rest of the national market, then p_j is a uniform price determined by interaction of demand and supply in the whole national market (the "one price"), $p_j = p^*$. Fig. 1 illustrates equilibria in local market of segment i. We assume that, apart from price, the only determinant of local demand is personal income per capita in the location. A change in personal income shifts local demand D along the quantity axis Q_i.

The left panel of the figure relates to the case when segments i and j are perfectly integrated. Because of the absence of impediments to trade between segments ("unrestricted commerce", as Cournot characterized it) the supply curve S in i is perfectly elastic. Let an increase in income shifts demand in segment i, say, from $D(M_i')$ to $D(M_i'')$, pushing the price in i up. However, this creates opportunities for arbitrage; then the good is exported from segment j to i. Additional supply due to this eliminates the rise in price in i. If a shift in local demand – e.g., from $D(M_i''')$ to $D(M_i'')$ – cause a decrease in price, the good is exported from i to j, so eliminating the difference in prices by reducing supply in i. (We neglect the time the arbitrage takes, considering arbitrage instantaneous.) Thus, local price in the integrated market does not depend on local demand.



(a) Perfectly integrated market (b) Segmented market

Fig. 1 Equilibria in spatial segment i of two-segment market

The right panel of Fig. 1 depicts the case of market segmentation. Arbitrage between i and j

needs transaction costs c_{ij} (as a percentage of the price of the good). When demand in segment i shifts from $D(M_i')$ to $D(M_i'')$, additional supply from segment j is not able to equalize the prices in i and j. It can reduce the price in i only to $p_i = p_j(1 + c_{ij})$, or $P_i = P_j + C_{ij}$, where $C_{ij} = \log(1 + c_{ij})$. When the arbitrage acts in opposite direction, $P_j = P_i + C_{ij}$, thus, $P_i = P_j - C_{ij}$. With $P_j - C_{ij} < P_i < P_i + C_{ij}$, the arbitrage does not work; local demand and local supply (S_L) determine price p_i. So, in a segmented market, in contrast to integrated markets, the price of the good depends on local demand (and eventually on income per capita).

Solving the equilibrium equation $q_D(p_i, m_i) - q_S(p_i,) = 0$, we can get p_i as a function of m_i (note that while the demand is local, the supply is not, being formed jointly by producers from all segments through arbitrage). Let $\beta = \frac{d\log(p_i)}{d\log(m_i)} = \frac{dP_i}{dM_i}$, then the above function can be represented as $p_i = p^* m_i^{\beta}$. The ratio of prices in i and j is $\frac{p_i}{p_j} = \frac{p^* m_i^{\beta}}{p^* m_j^{\beta}} = \left(\frac{m_i}{m_j}\right)^{\beta}$. If arbitragers from j sell the good in i, then $\frac{p_i}{p_j} = \left(\frac{m_i}{m_j}\right)^{\beta}(1 + \tau_{ij})$, where τ_{ij} stands for transportation costs as a percentage of the price of the good. We follow Engel and Rogers (1996) and many other authors by using distance to proxy transportation costs; we model them as $1 + \tau_{ij} = a l_{ij}^{\gamma}$, where l_{ij} stands for distance between i and j. In logarithms, we have the following equation:

$$P_{ij} = \alpha + \beta M_{ij} + \gamma L_{ij}, \qquad (1)$$

where $P_{ij} = \log(p_i/p_j)$ is the price differential, $M_{ij} = \log(m_i/m_j)$ is the income differential, $L_{ij} = \log(l_{ij})$, and $\alpha = \log(a)$.

If segments i and j are perfectly integrated, $\beta = 0$. This implies that the price differential is due to transportation costs only. Nonzero values of β evidence that market is segmented to some or other degree and serve as a measure of market segmentation (or, conversely, integration, implying that the smaller the β, the stronger the integration).

Taking account of random deviations (ε_{ij}), we obtain an econometric version of theoretical model (1):

$$P_{ij} = \alpha + \beta I_{ij} + \gamma L_{ij} + \varepsilon_{ij}. \qquad (2)$$

This regression is estimated over a set of all $N\times(N-1)/2$ segment pairs, N standing for the number of segments (cities in our empirical analysis). The expected values are: $\beta \geq 0$, $0 < \gamma < 1$ (the latter inequality suggests that that transportation costs rise with distance, however, with a decreasing rate). The units of distance measurement do not affect coefficient γ; only the regression constant, α,

depends on them. Note that it is a cross-sectional regression; it involves data for a given period. By estimating it separately for each available period, we obtain a temporal pattern of integration.

An important caveat is needed here. The price differential is antisymmetric with respect to the permutation of indices, $P_{ij} = -P_{ji}$, while distance is symmetric, $L_{ij} = L_{ji}$. Therefore, to ensure the positive dependence of the price differential on distance, the segment indices must be ordered so that all $P_{ij} \geq 0$.[4]

4 Data

The source of price data for the analysis is the Council for Community and Economic Research (C2ER), formerly the American Chamber of Commerce Researchers Association (ACCRA). Since 1968, this non-governmental institution publishes quarterly bulletin "Cost of Living Index" (formerly, "ACCRA Cost of Living Index"). Currently, it provides Excel files with data starting from 1990. Moeller (2020) briefly reviews this database.

The data include prices for more than 60 goods and services from over 300 cities of the US with a quarterly frequency. This rich data were and still are widely used by many researchers (some of them are mentioned in Section 2). In addition to prices for individual commodities, they are aggregated into six baskets: grocery items, housing, utilities, transportation, health care, and miscellaneous goods and services. The weights in the baskets are uniform across cities (they are drawn from the BLS Consumer Expenditure Survey); the costs of baskets are represented relative to the sample average. The commodities are homogeneous across cities as described in explicit instructions for data collectors in C2ER (2017).

The feature of these data is that their collection is voluntary. As a result, there are many irregularities in the dataset. The sample of cities varies from quarter to quarter; some new cities appear in the sample, while some others disappear. Moreover, even if a city presents in the sample permanently, it can contain a number of missing observations. Since this phenomenon is especially peculiar to the 1990s and earlier, we have limited our time span to 2001–2015.

We analyze the US market for an aggregated good, namely, grocery basket. The basket covers 27 foods; for its composition and weights, see C2ER (2017). As for the spatial sample, we have chosen cities that are permanently present in the dataset in 2001–2015 and have no more than two missing observation a year. We interpolate such missing observations as the arithmetic averages of

[4] Most likely, failure to meet this condition resulted in negative, albeit statistically significant, estimates of β and γ in Arapovic and Karkin (2015).

neighboring observations. Because of irregularities in the dataset we are not able to cover all states of the US. The resulting sample consists of 66 cities from 39 states.

The cities are: Birmingham, Huntsville, Montgomery (Alabama); Anchorage (Alaska); Phoenix, Tucson (Arizona); Little Rock-North Little Rock (Arkansas); Fresno, Los Angeles-Long Beach, San Diego, San Francisco (California); Colorado Springs, Denver (Colorado); Hartford (Connecticut); Wilmington (Delaware); Jacksonville, Miami-Dade County, Orlando, Tampa (Florida); Atlanta (Georgia); Boise (Idaho); Chicago, Springfield (Illinois); Cedar Rapids, Des Moines (Iowa); Lexington, Louisville (Kentucky); Baton Rouge, Lafayette (Louisiana); Baltimore (Maryland); Boston (Massachusetts); Grand Rapids (Michigan); Jackson (Mississippi); Kansas City, Springfield, St. Louis (Missouri); Omaha (Nebraska); Las Vegas, Reno-Sparks (Nevada); Las Cruces (New Mexico); New York-Manhattan (New York); Charlotte, Durham (Ohio); Cleveland (North Carolina); Oklahoma City, Tulsa (Oklahoma); Portland (Oregon); Philadelphia, Pittsburgh (Pennsylvania); Charleston-North Charleston, Columbia (South Carolina); Knoxville, Memphis (Tennessee); Austin, Dallas, El Paso, Houston, Lubbock, San Antonio (Texas); Salt Lake City (Utah); Burlington-Chittenden County (Vermont); Richmond (Virginia); Spokane, Tacoma (Washington); Charleston (West Virginia); Green Bay (Wisconsin).

Since personal income per capita are available with the annual frequency only, we aggregate the costs of basket to make them also annual.[5] To do so, we apply a simple arithmetic average over four quarters. (One more reason for limiting the spatial sample was that data on personal incomes per capita were not available for some cities.) The source of data on personal income per capita is the Bureau of Economic Analysis (BEA 2017).

The distance data have been obtained with the use of the Google online application Maps Distance Matrix API (Google 2017).

5 Empirical results

Table 1 reports descriptive statistics of variables. The statistics are computed over 2145 pairwise observations generated by 66 cities. As mentioned in Section 3, city indices *i* and *j* are arranged so that all $P_{ij} \geq 0$. Thus, P_{ij} is equivalent to the absolute price differential, providing a pattern of price dispersion in the city sample under consideration.

[5] In some ways, this aggregation can be considered an advantage, as it "absorbs" the time taken up by arbitrage.

Table 1 Descriptive statistics of variables

Year	Price differential			Income differential	
	Mean	Standard deviation	Maximum p_i/p_j	Mean	Standard deviation
2001	0.117	0.091	1.73	0.117	0.324
2002	0.133	0.101	1.83	0.119	0.311
2003	0.129	0.101	1.87	0.101	0.303
2004	0.147	0.111	1.79	0.109	0.304
2005	0.132	0.105	1.71	0.114	0.317
2006	0.127	0.107	1.81	0.129	0.332
2007	0.133	0.112	1.90	0.128	0.331
2008	0.118	0.099	1.72	0.095	0.330
2009	0.116	0.093	1.79	0.100	0.311
2010	0.119	0.101	1.89	0.099	0.310
2011	0.116	0.098	1.84	0.102	0.317
2012	0.108	0.093	1.83	0.100	0.332
2013	0.118	0.095	1.71	0.099	0.333
2014	0.099	0.079	1.60	0.115	0.340
2015	0.116	0.089	1.61	0.123	0.333

On average over 2001–2015, the mean price differential is 0.122. This suggests that the average difference in the costs of the grocery basket is 13% ($= (\exp(\bar{P}) - 1) \cdot 100$), changing between 10.4% in 2014 and 14.2% in 2007. The deviations from the mean are in the range of 108.2% (2014) to 117.4% (2004). The fourth column of the table reports the maximum price ratio across all city pairs (the minimum is 1 or near it by construction of the price differential). The price in the most expensive city exceeds the price in the cheapest city by no more than 1.9 times. This is peculiar to 2007 and 2010, while in 2014 and 2015 the maximum ratio decreases to about 1.6.

The dispersion of income differential is almost three times greater than that of price differential, while their means are comparable. However, the latter is due to the fact that the income differentials may be negative. Although prices in poorer locations are usually lower than in richer locations, the number of cases of the inverse relationship (i.e. negative income differential) in our sample is not small: from about one third to almost 40% depending on the year. The dispersion of incomes is much greater than that of prices. The maximum ratio of incomes, m_i/m_j, ranges from 4.17 to 5.12 for positive income differentials, and from 2.08 to 2.90 for negative ones.

The geometric mean of distances is about 1,700 km (1,050 miles); the arithmetic mean is 2,100 km (1,300 miles). The distances range from circa 50 km (30 miles) to 8,000 km (5,000 miles).

Table 2 presents results of estimation of regression (2) by year; the last line reports estimates of pooled panel regression. All estimates are statistically significant and have the expected positive

sign. Besides, the coefficient on distance is less than one as expected. Since the degree of market segmentation, β, is positive and significantly differs from zero in all cases, the market under consideration is not perfectly integrated. This is consisted with results of previous studies reviewed in Section 2.

Table 2 Estimation results

Year	Income differential (β)		Distance (γ)		R^2
2001	0.123	(0.007)	0.028	(0.002)	0.256
2002	0.141	(0.008)	0.023	(0.003)	0.224
2003	0.119	(0.008)	0.033	(0.003)	0.197
2004	0.132	(0.008)	0.042	(0.003)	0.217
2005	0.134	(0.007)	0.033	(0.003)	0.225
2006	0.148	(0.009)	0.036	(0.003)	0.275
2007	0.159	(0.009)	0.042	(0.003)	0.303
2008	0.126	(0.007)	0.040	(0.003)	0.276
2009	0.133	(0.008)	0.026	(0.002)	0.253
2010	0.148	(0.009)	0.027	(0.003)	0.263
2011	0.149	(0.008)	0.026	(0.003)	0.288
2012	0.154	(0.007)	0.021	(0.002)	0.345
2013	0.142	(0.006)	0.021	(0.002)	0.289
2014	0.116	(0.005)	0.021	(0.002)	0.296
2015	0.113	(0.006)	0.024	(0.002)	0.229
Pooled	0.136	(0.001)	0.029	(0.001)	0.253

White heteroscedasticity-consistent errors are in parentheses; all estimates are significant at the level of 0.1%.

As all variables are logarithms, the regression coefficients are elasticities of price differential vis-à-vis an independent variable. Averaging the estimates over 15 years, we get $\bar{\beta} = 0.135$ and $\bar{\gamma} = 0.030$. (Note that these figures are close to estimates of the pooled panel regression.) Thus, a 1-percent rise in the income differential increases the price differential by 0.135%. The effect of distance is by one order of magnitude weaker: a 1-percent increase in distance gives increases rise in the price differential by 0.03%.

The degree of segmentation, β, fluctuates over time with no visible regularity, having statistically insignificant trend. The least value, 0.113, is observed in 2015; the greatest β, 0.159, is peculiar to 2007. However, the 95% confidence intervals of most estimates overlap, with the exception of 2007 and 2012 (Fig. 2). Therefore, differences in most estimates may be attributed to random reasons. Alterative versions are also possible, though. The reasons may be some

macroeconomic events influencing the US consumer market. For instance, the impact of the financial crisis that started in 2007 may explain the peak in β in that year. In addition, differences in inflation between cities could result in deterioration or improvement of market integration in some years.

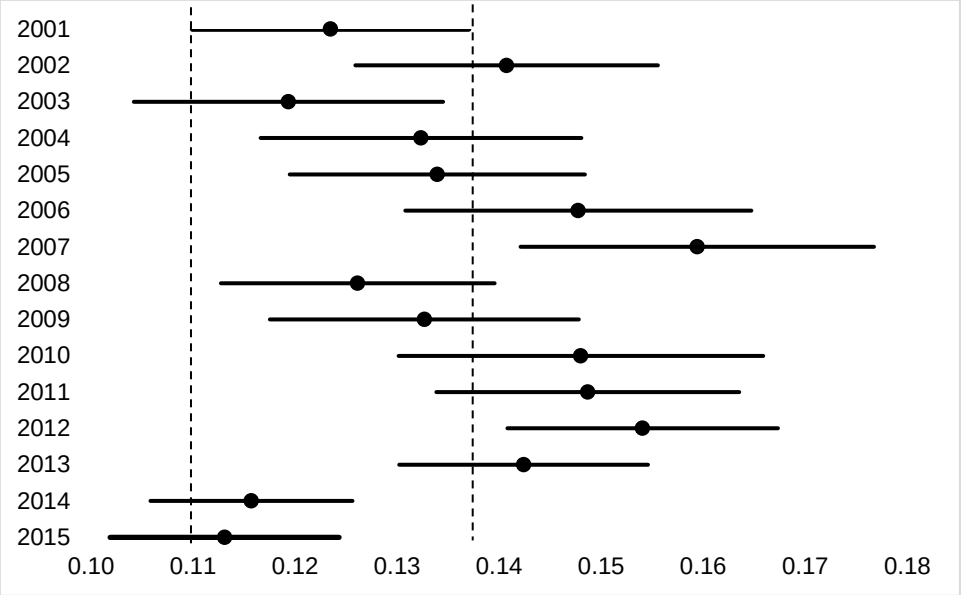


Fig 2 The 95% confidence intervals of estimates of β. Dash lines show the area of overlaps; dots are point estimates

Interestingly, the degree of segmentation, β, is strongly correlated with the maximum price differential, the correlation coefficient being 0.705. The correlation of β with the price dispersion measured by the standard deviation of price differential is moderate, amounting to 0.557, while the correlation with its mean is weak, 0.201.

The distance elasticity of the price differential, γ, fluctuates over time as well. It has the smallest values, 0.021, in 2012–2014, and the greatest values, 0.042, in 2004 and 2007. These fluctuations are not synchronous with fluctuations of β; the correlation coefficient of β and γ is 0.126.

Gluschenko and Karchevskaya (2010) obtained smaller values for the same market in 2000: β = 0.055 and γ = 0.003, the latter being statistically insignificant. Most probably, discrepancies with our estimates are due to small spatial sample used in that study.

The results presented in Table 2 leave open the question of whether the degree of market segmentation in the US is low or not. There are no theoretical or intuitive considerations that would

allow to judge this. Therefore, a comparison is needed with more or less similar studies for other countries. Arapovic and Karkin's (2015) results are not suitable for such a comparison because they relate to a small country (Bosnia and Herzegovina). We will use results for Russia from Gluschenko (2021), who also analyzes the cost of a food basket.

The empirical material for Russia differs from that for the US in the following ways. First, observations are regions, and not cities. The spatial sample covers 79 regions of Russia (3081 region pairs). Second, a staple basket is dealt with, composition of which is different from the composition of the grocery basket. The former includes 33 foods (recall that the US grocery basket includes 27 foods). Some foods in these baskets can be approximately matched with each other, while some foods are present in one basket and absent in another. Like the grocery basket, the staples basket has uniform weights across regions and time.

It is clear from the above that comparability of estimates between this study and study by Gluschenko (2021) is rather low. Therefore, the results of comparisons should be treated with caution.

We compare the degrees of market segmentation in the US and Russia graphically in Fig.3. The graphs are based on data from our Table 2 and Table 2 in Gluschenko (2021, p. 81).

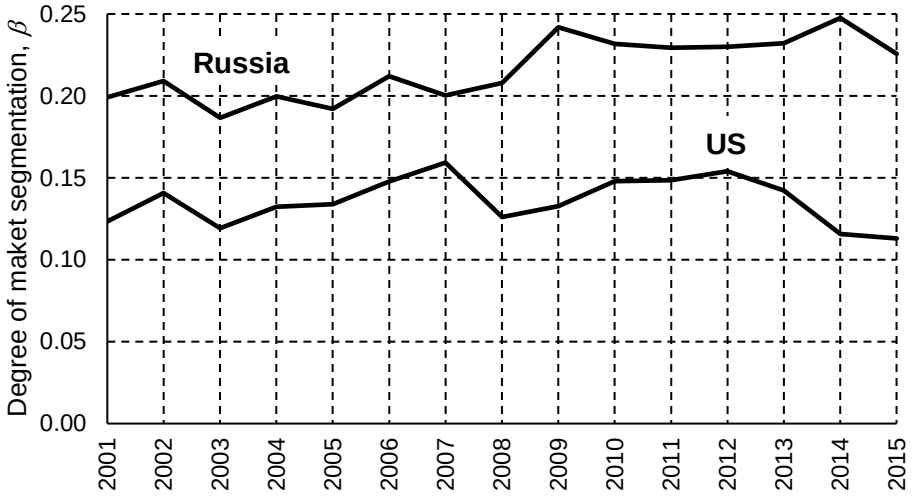


Fig 3 Comparison of market integration in the US and Russia

The graphs evidence that market integration in Russia is weaker that in the US. The average of β over 2001–2015 in Russia is 0.216; it changes between 0.187 (in 2003) and 0.248 (in 2014). This means that a 1-percent rise in the income differential increases the price differential on average by 0.216%. Leaving aside the differences in data, let us consider the possible reasons for the greater

segmentation of the Russian market.

One important reason seems to be much greater price dispersion in Russia. The mean over 2001–2015 of price differential in Russia is 0.184; annual means change in the range of 0.164 (in 2002) to 0.204 (in 2012). In terms of the average difference in the costs of the staples basket, these figures are 20.2%, 17.8%, and 22.6%, correspondingly. The standard deviation of price differential averaged over 2001–2015 is 0.201; the range of annual values is from 0.176 (in 2015) to 0.223 (in 2006). Thus, both the mean price differential and its standard deviation are about twice as much as in the US. On average over 2001–2015, the price in the most expensive region is 3.4 times higher than the price in the cheapest region, ranging from 2.8 in 2015 to 4 in 2006. (Recall that the maximum is 1.9 in the US.)

In turn, this may be due to Russia's larger territory and its geographical features. It is interesting to note that the global crisis affected different countries differently. It started in 2007 in the US, presumably manifested itself in the rise in market segmentation. This effect was short, though; the degree of segmentation fell in the next year. The crisis reached Russia in the end of 2008; its results became seen in 2009. As in the US, it contributed to weakening market integration. However, unlike the US, integration has not recovered in subsequent years; the value of β only slightly decreased in 2010–2013. This effect is entirely due to Russian difficult-to-access regions. There are several such regions (mostly in the north-east of the country) with low transport accessibility, arbitrage with which is almost (or entirely) impossible. Results for a subsample without these regions in Gluschenko (2021) do not indicate a significant impact of the crisis (moreover, the estimates of β are smaller in that subsample). The maximum prices are just observed in the difficult-to-access regions.

As regards distances (measured between capital cities of Russian regions), their geometric average is about 2,400 km (1,500 miles), and the arithmetic average is 3,600 km (2,250 miles). The maximum distance reaches 15,000 km (9,300 miles). Given longer distances than in our sample of the US cities, we can expect greater effect of distances in Russia. Fig. 4 depicts estimates of the coefficient on distance, γ, for the US and Russia. As expected, distances play a more significant role in Russia. The average value of γ over 2001–2015 in Russia is more than three times greater than in the US, amounting to 0.097 (thus, a 1-percent increase in distance increases the price differential by 0.097%). Its range is 0.075 (in 2002) to 0.112 (in 2006), twice as wide as the respective range in the US.

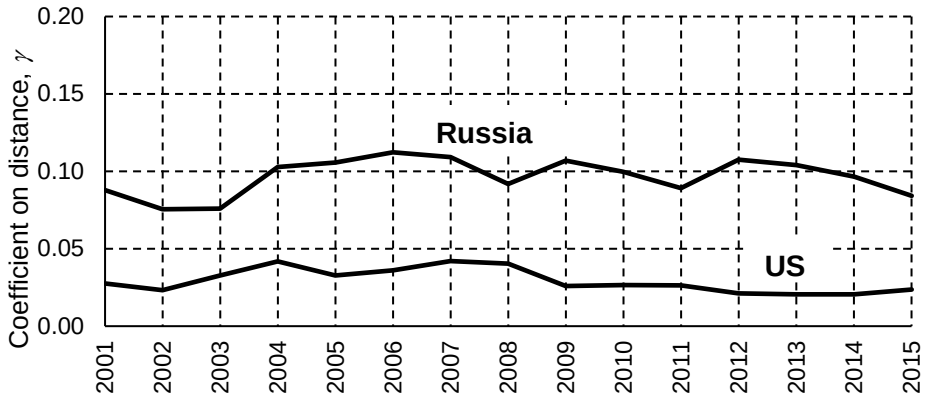


Fig. 4 Comparison of the distance effect in the US and Russia

The comparison does not allow us to draw a definitive conclusion about whether the degree of integration of the US market is high or low. The estimates for Russia have the same order of magnitude as the estimates for the US. Moreover, the differences between estimates can be explained by different data used and geographical differences. We can only hypothesize that the degree of market integration in the US may be deemed normal for a large country.

6 Conclusion

Based on the idea that dependence of local price on local demand (controlling for transportation costs) indicates a deviation from the law of one price, and its "strength" can measure the degree of market integration (segmentation), we estimate annual segmentation degrees of the US market for an aggregated good (grocery basket) over 15 years, from 2001 to 2015. The observations are 68 cities from 39 states of the US, generating 2145 city pairs.

The results suggest that the US market is not perfectly integrated; however, the integration degree of the US market is fairly stable over time. We also compare results for the US with results of a similar study for Russia. With a reservation that the empirical material is not fully comparable, we can conclude that the US market is integrated more strongly than the Russian market and that the integration degree in Russia is more volatile. The fact that the degree of market integration in Russia is comparable to that in the US by the order of magnitude is not surprising. By the 2000s, the Russian consumer market, in general, did not differ from markets in developed countries. Even taking account of different empirical material, it is reasonable to expect that market integration in Russia should be weaker than in the US because of longer distances between Russian regions that impede inter-regional arbitrage and the presence of a number of difficult-to-access regions.

A definitive conclusion about the strength of market integration in the US does not follow from the comparison of results for the US and Russia. We are able only to (hypothetically) deem the degree of market integration in the US normal for a large country. Similar studies covering other countries with large territories, such as Canada, China, Brazil, and Australia, could shed more light on this issue.

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