How Fast Is Convergence to the Law of One Price? Very (When Price Differentials Are Properly Measured)*

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Abstract

In a detailed product-level sample of monthly frequency, mostly tradable prices, the degree of persistence in price differentials is estimated. When the measure of cross-location price differential accounts for location specific effects, the median half-life of conditional convergence is about 4 months, significantly faster than in related studies of the law of one price. The equilibrium level of price differentials depends on the size of the main city in a county, but not on its geographical position.

Key words: Price Convergence, Law of One Price, Microeconomic Price Data

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1 INTRODUCTION

Why do large and persistent deviations from the Purchasing Power Parity, the identity of prices in two countries once converted to the same currency exist? In other words, what explains the large and persistent movements in the real exchange rates? One of the answers to these long-standing, still unresolved puzzles in international macroeconomics is the failure of the law of one price.

Being one of the most basic ideas in economic theory, the law of one price is defined by the combination of two statements. First, prices of dis-aggregated goods and services measured in the same currency should be identical across geographic locations in the long run. Second, temporary price misalignments should quickly disappear over time as consumers and firms can arbitrage away price differentials. While the idea is strongly rooted in appealing economic intuition, empirical analyses notoriously fail to establish support for it in the data; price differentials tend to remain persistent over time, fading away only relatively slowly.

There exist several potential explanations for the failure in the law of one price. These primarily include transportation costs, different currencies with sticky prices, labor market segmentation, tariff and non-tariff barriers to trade, fluctuations in nominal exchange rates, productivity differentials, pricing to market, aggregation bias etc. It is unlikely that any of these factors in isolation explains the slow mean reversion, it is more likely instead that some combination of them provides the answer. By focusing on the degree of persistence in relative price differentials across counties within a single country, the approach this paper adopts testing the law of one price plays down the importance of most of the above explanations. Indeed, the implied speed of convergence in this context is expected to be faster than in studies testing the absolute law of one price in international price data. An alternative interpretation of the law of one price is that equilibrium price differentials exist, but there is a relatively fast convergence to them over time. The time invariant component of price differentials may in turn depend on the relative economic development of locations and the distance between them.

The rest of the paper is organized as follows. The microeconomic price data are described in Section 2. Section 3 takes up measurement and specification issues. The results are presented in Section 4. For comparison, Section 5 briefly surveys results in the related literature. Conclusions follow in Section 6.

2 DATA

The empirical analysis builds on a data set of store level consumer prices recorded in Hungary.[1] The sample is drawn from the larger sample of consumer prices collected at the monthly frequency for the computation of the CPI by the Central Statistical Office, Hungary, and consists of cross-sections of price observations of twenty homogenous items, mostly specific food products and a few services. The particular products are

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[1] The sample is similar to and overlaps with the one utilized in Ratfai (2002). The current sample is longer but due to data availability it is restricted to a smaller number of products.

selected from the full CPI database with an eye to obtaining very narrowly defined (according to size, branding, type and flavor), continuously available items with negligible variation in non-price characteristics. The food items fall in the tradable (perishable or non-perishable) category, while the services clearly tend to be non-tradable. The products are the following: pork chops, spare ribs, pork leg, beef round, beef shoulder, pork liver, italian sausage, boiling sausage, carp, curd, lard, fat bacon, smoked boiled bacon, flour, sugar, dry biscuits, tomato paste, vinegar, car driving school, movie ticket.

Table I lists the products together with the expenditure weight attached to them in computing the aggregate CPI and their relative expenditure weight in the current sample. The table also reports the mean and standard deviation of product level monthly inflation rates. The standard deviation figures suggest that product-specific monthly inflation tends to be extremely variable. The figures also indicate that monthly inflation rates can substantially differ across items, their mean ranges from 1.19 to 2.82 percent, and their standard deviation from 1.64 to 8.06 percent.

The data are available over a period of 10 years (120 months), from 1992:1 until 2001:12 at the monthly frequency. In each month, there are 100-150 price observations (on average about 130) for each product. The number of stores is stable with an average standard deviation of less than 3 over time. The prices are recorded in 20 geographically dispersed locations including all the 19 counties in Hungary and the capital city, Budapest. Given its central economic and geographic importance in Hungary, Budapest is used a benchmark city in much of the analysis. Table II gives a complete list of locations. Although stores in the sample are identified by their location, they are not longitudinally matched. The set of stores visited is still likely to be stable over time as the data collectors are formally instructed by the CSO to keep the set of stores unchanged as much as possible.

County-product-time specific prices are computed as the arithmetic average of potentially different price observations of the same item, the same county, same month, recorded in different stores. Overall, the resulting panel of price indices has three dimensions with 20 counties, 20 products and 120 months. The panel of a total of 48,00 data points is balanced; i.e. there are no missing average price observations.

Despite the turbulent economic environment during economic transition, aggregate inflation throughout the sample period was relatively moderate in Hungary. Year-to-year change in the monthly aggregate CPI and its food component are plotted in Figure I. The graphs show that after an initial burst, annual aggregate inflation decelerated until by 1994. After reaching a minimum of about 15 percent on an annual basis, inflation dynamics eventually turned on an increasing path reaching about 30 percent at its peak in early 1995. Starting during the second quarter of 1995 shortly after the announcement and implementation of a macroeconomic adjustment package in March 1995 and lasting through the rest of the sample period, a disinflationary trend took effect.

3 MEASUREMENT AND SPECIFICATION

To measure the speed of price convergence across counties of Hungary, first, the product-specific price differential between two counties is defined as the log-difference in price levels, the benchmark and the actual ones. That is,

$$Q_{ij,k,t} = \ln(P_{i,k,t} / P_{j,k,t}),$$

where $P_{j,k,t}$ is the price of item k in county j in month t. Correspondingly, $P_{i,k,t}$ is the price the price of item k in month t observed in the benchmark location i, Budapest. In the absence of barriers to price equalization through arbitrage, the price differential measure, $Q_{ij,k,t}$, would be zero.

Persistent income differences and local non-tradable factors of production, transportation costs and other potentially time-invariant barriers of trade may create a constant wedge among retails prices at different counties. This reasoning implies that price differentials may never fully converge to zero, they just approach to a common, potentially non-zero level, being identical across different locations. To account for these time-invariant, county specific price differentials in the regression analysis[2], it is useful to make a transformation of price differentials and define their mean-difference as

$$Y_{ij,k,t} = Q_{ij,k,t} - \frac{1}{T}\sum_{t=1}^{T} Q_{ij,k,t} \,.$$

where the number of time periods $T = 120$.

The fundamental objective of this study is to characterize the degree of persistence in *intra*-country price differentials, after taking into account non-vanishing differences across locations. To do so, a series of panel unit tests are employed as developed by Levin, Lin and Chu (2002).[3] The main advantage of panel unit root tests over univariate ones is increased power, that is, the ability to reject the null of non-stationarity when it is actually absent. Given the definition of demeaned price differentials, the following basic regression specification is employed:

$$\Delta Y_{ij,k,t} = \beta Y_{ij,k,t-1} + \sum_{p=1}^{s(k)} \gamma_p \Delta Y_{ij,k,t-p} + \omega_{ij,k,t} \,,$$

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[2] See Parsley and Wei (1996) and Vidovic (2003). Vidovoc points to a possible bias in the speed of convergence when persistent cross-county differences in price levels are left unaccounted for. He suggests for demeaning the price differential data.

[3] An alternative unit root test in a panel setting that allows for individual heterogeneity is the one developed by Im, Pesaran and Shin (1996).

where $Y_{ij,k,t}$ is the price differential measure. The optimal lag structure in the regressions is determined by a series of product-county specific *t*-tests. As a result, the number of lags differs across counties. The parameter of primary interest in the empirical specification is β, capturing the degree of persistence in price differentials. The null of non-stationarity in price differentials is associated with the hypothesis of $\beta = 0$. The one-sided alternative is that the autoregressive coefficient is negative, $\beta < 0$. Levin, Lin and Chu demonstrate that the resulting *t*-statistic on the autoregressive coefficient is distributed normally.

Given the estimated autoregressive coefficients, the implied half-life of deviations from the law of one price are calculated under the assumption that the price differential process is AR(1). The reason for neglecting higher order terms in the impulse response is that, one the one hand, higher order persistence is rarely significant in the present application, so the associated terms can be safely ignored. On the other hand, higher order lags differ across counties that in turn makes it difficult to account for them in characterizing persistence at the product level. Overall, for each product, the half-life of deviations is obtained as $h = \ln(0.5)/\ln(1+\beta)$.

4 RESULTS

The estimated autoregressive coefficients from the panel unit root test of Levin, Lin and Chu (2002) along with the corresponding *t*-statistics are reported in the first two columns of Table III. Comparisons with the critical values provided by Levin, Lin and Chu suggest that the null of unit root can be safely rejected for all items. The point estimates for the root (one plus the estimate of the autoregressive parameter) are significantly different from one; indeed, they range from 0.73 to 0.94. The median value is 0.84. Serving as another metric on the degree of persistence in price differentials, the corresponding estimates for the half-life of divergence from the (relative) law of one price are displayed in the last column in Table III. The results suggest that price differentials are fading away fairly fast, the estimated half-lives are between 2.18 and 12.03 months, with a median value of 3.99 months. Price differentials disappear extremely fast, certainly faster than found anywhere else in previous related studies.

It is instructive to examine what product characteristics are associated with the degree of persistent in price differentials. First, the speed of price adjustment across geographical locations depends on the type of the product. Price differentials of services appear particularly persistent, while adjustment in food prices (perishable or non-perishable) takes place fast. Interestingly, prices of perishables adjust only slightly faster than prices of non-perishable; the half-life of their adjustment is 4.29 months compared to the similar figure of 4.37 months for non-perishable ones. Moreover, the relatively slower adjustment of non-perishables is really due to one outlier, tomato paste. Taking that item out, the average half-life gets reduced to 3.34 months. These results are puzzling, as one would expect non-perishable items to be more easily transportable and thus having less persistent price differentials.

Second, trend inflation and volatility in product prices also appear to be related to the degree of persistence. In particular, as shown in Table I, trend inflation is associated with the mean change in product-specific price levels, while volatility is defined as the standard deviation of product-specific price levels. The half-life of deviation from the law of one price and the trend inflation are positively correlated with a coefficient of 0.38. At the same time, the half-life of deviations is negatively correlated with the volatility measure with a coefficient of –0.18; though this latter correlation is statistically not significant. These observations suggest, consistently with the two-sided (S,s) pricing model of Tsiddon (1993) that a higher trend and less variability in price changes increase the probability of price adjustment, thus fastens the elimination of cross-location price differentials.

To assess the robustness of the above results, and to conform to some of the rest of the related literature, the panel unit root test described above are applied to price differentials not purged from time-invariant components. The results are summarized in Table IV. The first point to notice is that the price differential series are more persistent than in the demeaned price differential case. The autoregressive coefficients range from 0.83 to 0.99 with a median value of 0.94. Indeed, one of product price, movie price tickets series cannot be rejected to be non-stationary. The corresponding half-lives of unconditional convergence are also higher with the largest value being 3.65 months and lowest value being 115.18 months. The median half-life is however only 11.6 months, slightly shorter than one year. Service price differentials clearly disappear only very slowly, while the half-lives of deviations for perishable and non-perishable items are 6.63 and 5.59 months, respectively.

The speed of adjustment to the law of one price is ultimately shaped by the size and the degree of coordination of pricing shock and by the size of impediments to regional adjustment. In order to learn about the importance of arbitrage costs as proxied by the distance among geographical locations in time-invariant cross-county price differentials, the following cross-sectional gravity type equation is specified

$$\frac{1}{T}\sum_{t=1}^{T} Q_{ij,k,t} \equiv \overline{Q_{ij,k,t}} = a + b * d_{ij} + c * u_j + \varepsilon_j \,.$$

In the regression equation, $\overline{Q_{ij,k,t}}$ is the county-specific mean price differential relative to the benchmark location *i*, obtained by averaging over all products *k* and months *t* in county *j*. The variable d_{ij} is the distance between the main city in district *j* and the benchmark location, Budapest. u_j is the number of people living in county *j* to proxy for urbanization and development in the county. As expected, the regression results displayed in Table V show that a higher distance from the benchmark increases the mean price differential, though the relationship statistically is insignificant. At the same time, the population size of the county has a statistically significant negative impact on the mean price differential. This indicates that the constant component of price differentials tends to shrink as the difference in the level of development as proxied by population gets reduced.[4]

[4] Estimating specifications with quadratic terms for distance and populations show the absence of statistically significant non-linear effects.

5 RELATED LITERATURE

The literature on convergence to the law of one price has primarily focused on cross-country and aggregate level price differentials. The consensus view in the international context on the estimated half-life of deviations from purchasing power parity has been three to five years; that is, convergence is present but it is very slow.[5] The current study differs from much of the PPP literature in two fundamental dimensions. First, it examines price convergence among locations that are geographically close to each other and that share the same currency. Second, it uses a sample of highly dis-aggregated, product level prices of very narrowly defined homogenous items. A number of studies have particularly close bearings on the current analysis.

First, Parsley and Wei (1996) estimate the half-life of deviations of prices across U.S. cities. The price data used is at the quarterly frequency from the first quarter of 1975 to the fourth quarter of 1992, and consist of prices of 51 products observed in 48 locations. The products are divided into nonperishable, perishable and services categories. The main variable of interest is the log-difference in price levels, without demeaning between individual cities and the benchmark city, New Orleans. First, they reject the null hypothesis of random walk for most items. They conclude that the estimated median half-life for nonperishable goods is 5.28 quarters, for perishable goods

[5] See Rogoff (1996). For instance, Frankel (1990) estimated the half-life of PPP deviations to be 4.6 years, Wei and Parsley (1995) found it to be 4 to 5 years, Lothian and Taylor (1996) 4.7 years, Abuaf and Jorion (1990) 3.3 years.

4.05 quarters, and for services 15.4 quarters. To test for non-linear effects, Parsley and Wei add a quadratic term as well, and they find that higher price differential is closed at a faster rate than a smaller price differential. Finally, they also analyze the dependence of the mean absolute deviation in and the variability of price-differentials on distance. They find positive association in both specifications, implying increasing costs to arbitrage as two cities being further apart from each other.

Building on the analysis of Parsley and Wei, Cecchetti, Nelson and Sonora (1999) study price convergence in a 78 years long panel of annual price indices in 19 US cities. Using a series of panel unit-root tests, they find a surprisingly low speed of convergence, with a half-life of about 9 years. Moreover, they are unable to reject the null of non-stationarity in price differentials when examined in univariate a unit-root tests. The degree of persistence in monthly frequency price differentials found in the present paper is smaller than the related quarterly frequency findings of Parsley and Wei, or especially the annual frequency findings of Cecchetti, Nelson and Sonora.

Engel and Rogers (1994) study the *intra*-country versus *inter*-country implications of the convergence to the law of one price. The specific objective is to quantify the effect the U.S. – Canadian border has on the price differentials between U.S. and Canadian cities, as compared to price differentials between cities in one of the two countries. Engel and Rogers use dis-aggregated, bimonthly consumer price indices of 14 commodities in 23 U.S. and Canadian cities, transformed into log-differences price differentials between two cities. The measure of volatility for each the 228 pair of commodity price indices is defined as the standard deviation of relative prices. The approach adopted by Engel and Rogers provides a particularly appealing tests of the law of one price as within-country real exchange rate fluctuations are unaffected by nominal exchange rate shocks.

In a gravity type, cross-sectional empirical specification, they demonstrate that the border is a strong determinant of relative price volatility, once distance and city effects are controlled for. Engel and Rogers compute a counterfactual border effect of two cities lying just across the border appearing to be 75,000 miles apart. The message of the analysis is clear: the currency in which price are denominated is important in itself and prices across the border tend to differ more than prices within a country for a given distance. The explanation they offer and find some evidence for is that sticky prices set in local currencies generate the excess volatility in price differentials.

In a closely related study examining price differentials in the US and Mexico, Rogers and Smith (2001) arrive at similar results.[6] Among other things, they find that the border effect is present across US-Mexican city pairs, but does not fall significantly when examined only in cities lying very close to the geographical border. This result confirms that the 'border effect' is not strictly a geographical phenomenon, rather it may be related to other features, such as the tax system or the currency regime of national economies.

Finally, Imbs *et al* (2002) analyze the possible bias in aggregate price data due to differences in the dynamics of dis-aggregated price data comprising of the aggregate ones. They show that failing to account for heterogeneity across product categories for realistic price dynamics produces downward-biased estimates of the speed of convergence, leading to upward-biased estimates of the half-life of price differentials. They demonstrate that the bias can be economically substantial: the estimated half-life is

[6] See also Wolf (2000) and Beck and Weber (2001).

39 months in a fixed effect panel model using price data in the European Union, while the same figure is 27 months in more dis-aggregated, sector level data. The results in general suggest that the appropriate testing ground for the study of price differentials are in microeconomic price data, as pursued in the current study.

6 CONCLUSION

This study analyses the *intra*-country convergence of price differentials using a long, monthly frequency panel of prices of highly dis-aggregated items in Hungary. The advantages of using microeconomic price data from the same country are manifold in this context. In particular, it mitigates the role of exchange rate fluctuations, factor and good market separation and barriers to trade in accounting for the speed of convergence in price differentials across different geographic locations. In addition, the use of micro level price data allows for controlling aggregation effects potentially biasing inference on potential sources of non-stationarity.

In contrast to a large portion of the literature on Purchasing Power Parity and the law of one price, the findings here strongly reject the null hypothesis of price differentials being non-stationary. Indeed, the implied half-lives in general show very fast convergence in prices. The speed is slower for non-tradable services than for tradable food items. If interpreted as fast convergence to a potentially non-zero time-invariant price differential, the law of one price appears to rule. The time-invariant price differential in turn depends on the absolute size of the geographic locations as proxied by the population in the main city in the county, but on the distance of the main city from the benchmark location.

The degree of persistence in price differentials has crucial implications for inflation dynamics in countries of joining the EU and eventually the EMU. The major potential tension is between the economic and monetary integration of currently low-price accession countries bringing about price level convergence and the need of keeping inflation at levels appropriate for monetary integration in these countries. The true quantitative consequences of this tension are determined in turn by the extent to which inflation differences are explained by forces other than convergence in price levels.

References

Abuaf, N. and P. Jorion (1990): Purchasing Power Parity in the Long Run, *Journal of Finance*, pp. 157-174

Beck, G. W. and A. A. Weber (2001): How Wide Are European Borders? New Evidence on the Integration Effects of Monetary Unions, *manuscript*

Cecchetti, S., N. Mark and R. Sonora (1999): Price Level Convergence Among United States Cities: Lessons for the European Central Bank, *manuscript*

Engel, Ch. and J. Rogers (1996): How wide is the border?, *American Economic Review*, pp. 1112-1125

Engel, Ch. and J. Rogers (1999): Violating the Law of One Price: Should We Make a Federal Case out of It?, NBER Working Paper #7242

Im, K. S., H. Pesaran and Y. Shin (1997): Testing for Unit Roots in Heterogeneous Panels, *manuscript*

Imbs, J., H. Mumtaz, O. Ravn and H. Rey (2002): PPP strikes back: Aggregation and the Real Exchange Rate, NBER Working Paper #9372

Levin, A., C. Lin and C. Chu (2002): Unit Root Tests in Panel Data: Asymptotic and Finite-Sample Properties, *Journal of Econometrics*, pp. 1-24

Lothian, J. and M. Taylor (1996): Real Exchange Rate Behavior: The Recent Float from the Perspective of the Past Two Centuries, *Journal of Political Economy*, pp. 488-509

Parsley, D. and S-J. Wei (1996): Convergence to the Law of One Price without Trade Barriers or Currency Fluctuations, *Quarterly Journal of Economics*, pp.1211-36

Ratfai, A. (2002): Inflation and Relative Price Asymmetry, *manuscript*

Rogers, J. and H. Smith (2001): Border Effects within the NAFTA Countries, *Board of Governors, International Finance Discussion Papers #698*

Rogoff, K. (1996): The purchasing power parity puzzle, *Journal of Economic Literature*, pp. 647-668

Sonora, R. (2002): Bivariate Relative City Price Convergence in the United States: 1918-1997, *manuscript*

Tsiddon, D. (1993): The (Mis)Behavior of the Aggregate Price Level*, Review of Economic Studies*, pp. 889-902

Vidovic, S. (2003): Convergence to the Law of One Price, *manuscript*

Wolf, H. (2000): (Why) Do Prices Differ Across US Cities? Balassa-Samuelson versus 42nd Street, *manuscript*

FIGURE I

Annual CPI Inflation in Hungary, Monthly Data

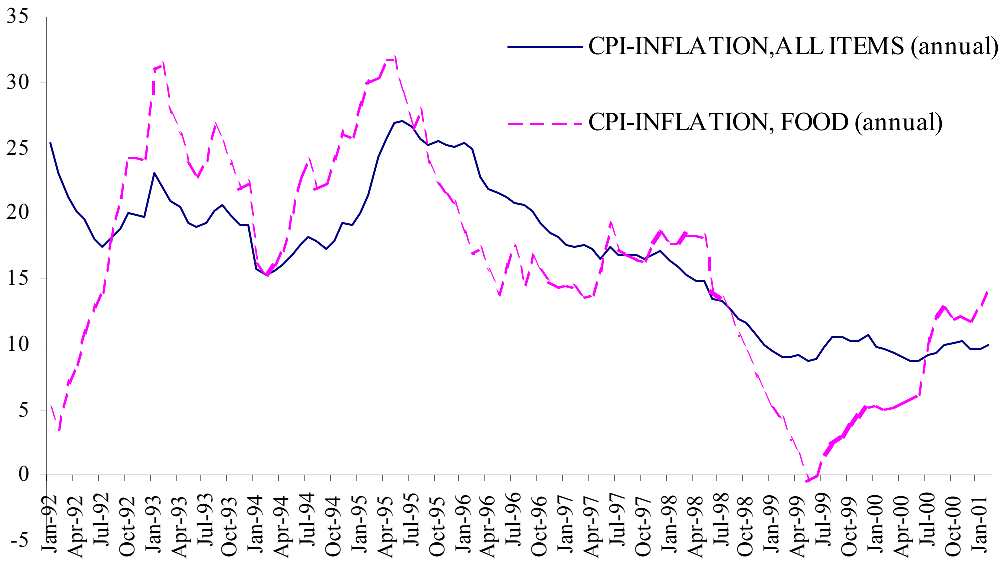

TABLE I
PRODUCTS IN THE SAMPLE

Product Code	Product Name	Type of Product	Absolute Weight	Relative Weight	Mean	Standard Deviation
10001	Pork, Chops	p	0.49	13.65	1.22	5.09
10002	Spare Ribs, with Bone	p	0.19	5.29	1.55	5.21
10003	Pork, Leg without Bone and Hoof	p	0.77	21.45	1.18	5.36
10102	Beef, Round	p	0.04	1.11	1.83	2.38
10103	Beef, Shoulder with Bone	p	0.04	1.11	1.97	2.46
10301	Pork Liver	p	0.12	3.34	1.75	2.86
10603	Sausage, Italian Type	p	0.17	4.74	1.71	2.88
10605	Sausage, Boiling	p	0.17	4.74	1.80	2.91
10801	Carp, Living	p	0.06	1.67	1.93	1.87
11302	Curd, 250g	p	0.16	4.46	2.33	3.67
12101	Lard, Pork	n	0.13	3.62	2.10	8.06
12201	Fat Bacon	n	0.07	1.95	2.81	3.85
12203	Smoked Boiled Bacon	n	0.07	1.95	2.62	3.90
13002	Flour, Prime Quality	n	0.28	7.80	1.98	3.35
13501	Sugar, White, Granulated	n	0.53	14.76	1.59	2.55
13801	Dry Biscuits, without Butter, Packed	n	0.05	1.39	1.63	1.86
14424	Tomato Paste	n	0.03	0.84	1.41	1.91
15208	Vinegar, 10 hydrate	n	0.05	1.39	1.19	2.99
66105	Car Driving School, Full Course	s	0.16	4.46	2.58	7.85
66301	Movie Ticket, Evening, 1-6 Rows	s	0.01	0.28	2.19	1.64
			3.59	100.00		

Notes 1 Figures are compiled from various consumer price statistic booklets of the Central Statistical Office, Hungary.
2 Products are narrowly defined according to size, branding, type and flavor.
3 Weights are expenditure-based. *Absolute Weight* is taken from the 1995 CPI. *Relative Weight* reflects weight of the item in this particular sample.
4 *Mean* is the average of monthly inflation rates. *Standard Deviation* is the standard deviation of monthly inflation rates.
5 The type of product can be perishable food (*p*), non-perishable food (*n*) and service (*s*). The classification is judgmental; it is based on a subjective examination of product characteristics.

TABLE II
LOCATIONS IN THE SAMPLE

Code	County / Main City	Population	Distance to Budapest
2	Baranya / Pécs	405 000	197
3	Bács-Kiskun / Kecskemét	545 000	83
4	Békés / Békéscsaba	396 000	203
5	Borsod-Abaúj-Zemplén / Miskolc	745 000	178
6	Csongrád / Szeged	427 000	169
7	Fejér / Székesfehérvár	428 000	65
8	Győr-Moson-Sopron / Győr	436 000	124
9	Hajdú-Bihar / Debrecen	552 000	226
10	Heves / Eger	325 000	127
11	Komárom-Esztergom / Tatabánya	317 000	57
12	Nógrád / Salgótarján	219 000	112
13	Pest / Cegléd	1 106 000	70
14	Somogy / Kaposvár	336 000	188
15	Szabolcs-Szatmár-Bereg / Nyíregyháza	586 000	245
16	Jász-Nagykun-Szolnok / Szolnok	416 000	98
17	Tolna / Szekszárd	249 000	142
18	Vas / Szombathely	267 000	221
19	Veszprém / Veszprém	374 000	109
20	Zala / Zalaegerszeg	298 000	227
1	Budapest / Budapest	1 725 000	0

Notes 1 The county codes and the population data are compiled from the appropriate booklets of the Central Statistical Office, Hungary.
2 The distance figures represent the distance between the main city in the county and Budapest, taken from the official roadmap of Hungary.

TABLE III
SPEED OF CONVERGENCE – BASELINE

Product Code	Product Name	Type of Product	Coefficient	*t* value	Half-Life
10001	Pork, Chops	p	-0.267	-12.04	2.23
10002	Spare Ribs, with Bone	p	-0.232	-12.94	2.63
10003	Pork, Leg without Bone and Hoof	p	-0.256	-13.27	2.34
10102	Beef, Round	p	-0.132	-10.51	4.90
10103	Beef, Shoulder with Bone	p	-0.096	-8.81	6.87
10301	Pork Liver	p	-0.088	-7.67	7.52
10603	Sausage, Italian Type	p	-0.177	-10.82	3.56
10605	Sausage, Boiling	p	-0.118	-7.81	5.52
10801	Carp, Living	p	-0.213	-12.86	2.89
11302	Curd, 250g	p	-0.111	-9.50	5.89
12101	Lard, Pork	p	-0.272	-13.99	2.18
12201	Fat Bacon	p	-0.130	-8.59	4.98
12203	Smoked Boiled Bacon	p	-0.151	-9.37	4.23
13002	Flour, Prime Quality	n	-0.169	-10.06	3.74
13501	Sugar, White, Granulated	n	-0.219	-12.19	2.80
13801	Dry Biscuits, without Butter, Packed	n	-0.197	-11.12	3.16
14424	Tomato Paste	n	-0.078	-6.97	8.54
15208	Vinegar, 10 hydrate	n	-0.173	-9.70	3.65
66105	Car Driving School, Full Course	s	-0.068	-8.53	9.84
66301	Movie Ticket, Evening, 1-6 Rows	s	-0.056	-8.05	12.03
	Mean		-0.16		4.98
	Median		-0.16		3.99
	Standard Deviation		0.07		2.75

Notes: See Table I.

TABLE IV
SPEED OF CONVERGENCE – ROBUSTNESS

Product Code	Product Name	Type of Product	Coefficient	*t* value	Half-Life
10001	Pork, Chops	p	-0.061	-5.07	11.01
10002	Spare Ribs, with Bone	p	-0.152	-9.81	4.20
10003	Pork, Leg without Bone and Hoof	p	-0.080	-7.42	8.31
10102	Beef, Round	p	-0.024	-4.02	28.53
10103	Beef, Shoulder with Bone	p	-0.058	-6.57	11.60
10301	Pork Liver	p	-0.037	-4.82	18.38
10603	Sausage, Italian Type	p	-0.058	-5.86	11.60
10605	Sausage, Boiling	p	-0.057	-4.81	11.81
10801	Carp, Living	p	-0.045	-4.56	15.05
11302	Curd, 250g	p	-0.083	-7.99	8.00
12101	Lard, Pork	p	-0.173	-10.78	3.65
12201	Fat Bacon	p	-0.040	-5.84	16.98
12203	Smoked Boiled Bacon	p	-0.086	-6.79	7.71
13002	Flour, Prime Quality	n	-0.046	-5.35	14.72
13501	Sugar, White, Granulated	n	-0.145	-9.87	4.42
13801	Dry Biscuits, without Butter, Packed	n	-0.079	-6.25	8.42
14424	Tomato Paste	n	-0.036	-4.64	18.91
15208	Vinegar, 10 hydrate	n	-0.060	-5.87	11.20
66105	Car Driving School, Full Course	s	-0.022	-4.25	31.16
66301	Movie Ticket, Evening, 1-6 Rows	s	-0.006	-2.52	115.18
	Mean		-0.067		18.04
	Median		-0.058		11.60
	Standard Deviation		0.044		23.99

Notes: See Table I.

TABLE V
DISTANCE AND POPULATION REGRESSION

Variable	Coefficient Estimate	*t* statistic	*p* value
Constant	-0.03	-1.75	0.10
Population	-0.0412	-2.12	0.05
Distance	0.0231	13.27	0.73
R^2	0.23		
F test	2.45		

Notes 1 Distance is measured in 100 km.
2 Population is measured in million people.
3 The dependent variable is the average of demeaned price differentials.