BEHAVIOR OF RETAIL PRICES: A NOTE ON MARKET INTEGRATION IN THE U.S.

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International economists have been attacking the relevance of the law-of-one-price for several decades and some interesting essays testing this law in a foreign trade context are available. Nevertheless, it is generally assumed that this law prevails in domestic markets and, as a result, it has not been seriously tested in the context of retail trade. Such a task would be useful, not just to learn whether "markets are doing their job" but also to gain perspective on the rate of change of domestic market integration.

The purpose of this essay is to test a dynamic version of the law-of-one-price in U.S. retail markets where transportation costs have been falling. I use a sample of average annual prices over five U.S. cities for each of 17 groups of goods and services for a 43-year period. For each product group I test whether the coefficient of variation over this set of five average prices increased or decreased over time. The basic conclusion is that the impact of the law-of-one-price is not evident. This suggests that in the post World War II era, U.S. retail markets did not become more integrated, at least if integration is defined in terms of price convergence.

According to a static version of the law-of-one-price, in a perfectly competitive market the prices of a given good at different points in space tend toward equality. Any variations in these prices can be traced to factors such as differential transportation costs of goods to various locations in the market, the difficulties of learning about prices in different outlets (information costs), or differences arising from bulk purchasing, outlet location and hours, and store rents. Of these, only transportation costs and store rents should make any significant difference in explaining differences in average prices over cities; the other factors wash out because the unit of analysis is large.

A dynamic version of this law focuses attention on decreasing price dispersion accompanying a fall in transportation costs. Charles Kindleberger [1989, 66ff.] traces this common sense observation back to Adam Smith. It is generally assumed that a declining price dispersion over time indicates an increasing market integration.

PRICE DISPERSION ACROSS SPACE: THE ISSUES AT STAKE

The key element in testing a dynamic version of the law-of-one-price is the determination that the share of transportation costs in total costs has declined since 1950. Unfortunately, direct data to test this are not available and we must rely on indirect evidence.

Producer prices of transportation services rose from 1950 through 1990 at an annual rate of 3.2 percent.[1] This was considerably lower than the increase of the consumer price index (CPI), and it was also lower than for any commodity group under consideration in the statistical analysis below, except clothing. Such data on relative price trends say nothing, of course, about changes in cost shares. But if we make the reasonable assumption that the price elasticity of transportation expenditures for producers and retailers is less than unity, then the share of transportation costs, measured in current prices, should have declined. Although I was unable to locate data on the changing share of transportation costs in total retail sales to justify this assumption, indirect evidence from foreign trade accounts and also national income accounts supports this.[2] With the exception of store rents (discussed below), it does not seem likely that the other factors underlying variations in prices across the five large cities used in the sample have changed enough to influence trends in price variations. Thus, falling transportation costs are the key.

Five reasons can be offered to explain why the dynamic version of the law-of-one-price might not apply in the domestic economy at the retail level when comparing average prices of a particular commodity group across large cities.

Rising Search Costs

The narrowing of price differences across space assumes that consumers shop assiduously for the lowest price with little regard to the cost of search. But analysis of the economics of information from Stigler [1961] to the present, suggests that this is not always the case. For a relatively expensive and homogeneous commodity or service, an extensive price search is rational since such a search is relatively inexpensive and the potential savings are great. For instance, some consumers wishing to buy a new automobile are willing to search a considerable distance from their home. Comparisons are easy since the product is homogeneous and expected absolute savings are large. If, however, the good or service is heterogeneous, it takes longer to learn about and compare its various features. Similarly, if the good is relatively inexpensive, it is not worth spending much time shopping for the lowest price unless the good is bought repeatedly. In the course of economic growth, the value of shoppers' time rises since the value of alternative uses of time increases with a rising per capita income. Thus, as per capita income rises we would expect less search by consumers, especially for heterogeneous goods. As a result, we would also expect price dispersion over space to increase, not diminish.

Technological change can lower search cost. Several market innovations such as the increasing importance of catalogue sales [Oi, 1988] have in fact, lowered the cost of search, strengthening the forces of price convergence. Up to now, however, these innovations do not appear to have had a great impact on price convergence between cities, at least for the product groups examined below. Computer shopping, which has the same effect, is still in its infancy, and its impact remains to be seen.

Declining Arbitrage Possibilities

Price dispersion should be less for those goods or services for which arbitrage is easy, so that people can buy where the price is low and sell where the price is high. But arbitrage is difficult on the retail level for several classes of goods and services: those that have relative high selling costs, for instance highly heterogeneous goods; and those that cannot be easily transported, for instance housing and many personal services. With rising per capita and discretionary incomes, consumer purchases move away from standardized to more heterogeneous goods, or from relatively simple models of a commodity to more complicated and higher quality versions. On the consumer level, therefore, arbitrage becomes increasingly more difficult between cities and, as a result, price dispersion should become greater for product groups.

Changes in Local and State Government Actions

At a single point in time, price dispersion across different governmental units should be greater for those goods and services for which some carry a significant share of sales taxes. It is unclear, however, whether sales tax rates in different states have converged or diverged over time. The same might be said for other types of governmental actions, for instance, the setting of public utility prices. From the Bureau of Labor Statistics data set that I employed, it was not possible to take this factor systematically into account; nevertheless, it seems unlikely that such actions would have markedly influenced price dispersion across the five cities.

Changes in Rents of Retail Outlets

In the long run, land rents of retail outlets are incorporated into retail prices. If the divergence of these rents across the five cities changes over time, the divergence of retail prices should change in the same direction, other things remaining equal. It is difficult to generalize about these matters from first principles because several factors must be considered. Unfortunately, the empirical evidence is also ambiguous. Although I have no direct evidence on retail rents, I do have data on the relative prices of all rental units, as well as shelter costs including rents, household insurance, and similar expenses. Whereas on the one hand, the dispersion of the shelter costs reveals no secular trend, the dispersion of house rents alone has increased. With such mixed evidence, it is difficult to draw a firm conclusion. Nevertheless, the broader concept of shelter expenses seems to be more relevant for considering retail expenses than rents alone since the former include utilities and taxes.

The shift of retail sales from high-rent central-city areas to low- and medium-rent shopping malls in the suburbs creates additional problems. I make the reasonable assumption that this change occurred at roughly the same rate in all five cities so that it does not have a marked impact on price dispersion.

Changes in the Degree of Imperfect Competition

If the sale of certain goods on the retail markets is dominated by oligopolies or monopolies in particular localities, and if the degree of imperfect local competition changes over time, price dispersions over space can be affected. Unfortunately, little information is available on these matters. To sidestep the problem, however, I have confined myself to large cities where retail market imperfections are presumably not great enough to influence the comparisons.

Given the existence of rising search costs, declining arbitrage possibilities, and changes in rents, governmental actions, and imperfect competition, it is impossible to determine theoretically which is stronger: the law-of-one-price or these counteracting forces. It is necessary, therefore, to turn to the data to decide which set of forces is dominant. Such medicine, however, turns out to be easier to prescribe than to take. The available data are not totally suited since they are aggregates, rather than series for individual goods. Furthermore, it proved impossible to enter the theoretical factors listed above into the empirical analysis in a systematic fashion. At the present time all we can do is to investigate the empirical question of whether the dynamic version of the law-of-one-price holds, that is, whether price converged or diverged over space as relative transportation costs fell.

THE DATA

The root of the empirical problem facing anyone wishing to test the law-of-one-price is the difficulty in obtaining comparable retail price data for particular goods and services in different cities. The BLS collects sample price information for different goods and services for every city each month using a probability sample to calculate the CPI. As a result, the highly detailed data underlying the CPI must be transformed in some appropriate manner.

A recent and interesting data set that can be used for these purposes comes from a BLS project investigating absolute differences in prices of commodity groups over various cities and regional areas [Kokoski, 1991; Kokoski, Cardiff, and Moulton, 1992; and Primont and Kokoski, 1991]. Although data for only a limited number of product groups are available, they have been carefully adjusted to take quality differences into account and allow us to gain some idea of price differences both over various cities around the country and, in some cases, within metropolitan areas. For instance, within metropolitan areas prices seem similar for such categories as food, as we would expect. For other categories—particularly highly heterogeneous goods—prices can be quite different. For instance, for household furnishings and operations, retail prices in the Connecticut suburbs run about 60 percent of the prices paid in New York City.[3]

These BLS price comparisons on which I base my calculations come with a disclaimer: Since the calculations were made for an experimental program and no standard errors could be calculated, we cannot tell whether the differences are statistically significant.[4] As a result, any conclusions drawn from a statistical analysis of these data must be considered tentative. Nevertheless, the attempt seems worthwhile since the economic implications of price dispersion over space, and also the associated methodological problems in making the empirical tests, have received little attention by economists.

Such comparable price data in different cities are, unfortunately, only available for product groups at a single point in time, namely a 12-month period in the late 1980s. Nevertheless, they can be combined with monthly current price indexes for the relevant product groups for each of the five cities, so that price dispersion in a given product group can be measured over different localities for a long period. Although the weighting of the goods within each aggregate varied over time, it changed in the same way for each city so that the results should not be biased on this account. Further, my procedure is based on the reasonable assumption that the differences in average quality of the goods in the five large cities remained roughly the same over time. That is, although New York and Philadelphia bagels differ in price and certainly quality, the ratio of their qualities should not greatly change over time. For each of the cities, the average prices for a particular good are calculated from a number of different outlets; over time, therefore, the choice of outlet should also not make much difference. Finally, the statistical test is based on a key assumption about the quality of these time-series discussed in the Appendix.

SELECTION OF THE SAMPLE AND SOME PROBLEMS OF TESTING

For the sample I chose five cities for which the time-series indexes are almost complete for all product groups for the 43-year period: New York, Philadelphia, Detroit, Chicago, and Los Angeles. The choice of particular groups of goods and services to be examined raises some problems. I have tested all 17 series that were available, but only eight can be considered "true tests" of the law-of-one-price (these are so designated in the table). The remaining 9 either feature considerable governmental interference in the price (for instance, alcoholic beverages or motor fuels) or represent goods or services with presumably only a local market area (for instance, housing rents or medical services).

Since prices of all commodity groups were rising, the prices were standardized. The basic statistic used below is the unweighted coefficient of variation of prices (*CV*: the standard deviation divided by the mean) using the five data points for a given period. To avoid problems of seasonal corrections, I averaged the monthly statistics for each year before calculating the annual *CV*s. The statistical test is whether the *CV* series increased or decreased over time.[5]

A final problem arises from the nature of shocks to the system. In looking at graphs of the *CV*s over time, most series appear to have a trend. It is also clear that the calculated trend coefficient is highly sensitive to the interval selected. In addition, most of the series feature irregular up-and-down movements lasting for a number of periods that differ among the various series. This suggests that random shocks are persistent and that some of the series might be described best as random walks. In the most simple case, this means that the dispersion value in period t is equal to its value

in period $t-1$ plus a random factor so that the series could drift up or down over long periods. It is, of course, also possible to have a random walk around a rising trend. Fortunately, some statistical procedures are available to deal with this problem. I used a Dickey-Fuller test to determine whether the *CV*s can be characterized by a random walk.[6] It turns out that a random walk occurred in only a few cases, but such a regression showed the presence of random shocks with effects lasting for some years. This same regression also permits us to determine a trend value of the series of *CV*s taking such shocks into account (the "corrected" trend value in the table).

THE RESULTS

Table 1 reports some descriptive statistics for the 17 commodity groups for short and long periods. Trend values of the changes in the *CV*s, both uncorrected and corrected for the impact of long-lasting random shocks are presented. The results can be quickly summarized.

Average Annual Price Increases

For the most part the product groups in the sample average annual prices increased somewhat less rapidly than the entire CPI (1950 to 1993: 4.28 percent; 1976 to 1993: 5.70 percent).

Average Spatial Variation

For the most part differences in the price dispersion of various product groups accord with expectations. Clothing and household furniture, which are much more heterogeneous than foods, both exhibit a much greater price variation over space than food. Given the quite different supply and demand conditions for housing in different cities, the price variations over space are also high. Services also show considerable spatial price variation, in some cases because they can not be easily arbitraged (for instance, medical and entertainment services), in others because of the influence of government taxes or price-setting regulations (for instance, piped gas and electricity).

Simple Trend Calculation

From the first year to 1993, 36 percent of the eleven product groups show a statistically significant divergence of prices over space (a positive coefficient), while 45 percent reveal significant convergence (a negative coefficient). For the five "true test" product groups, these percentages are 40 and 60 percent.

From 1976 to 1993, 53 percent of all 17 product groups had statistically significant divergent prices and 35 percent, convergent prices. For the eight "true test" groups, these percentages are 43 and 57 percent.

TABLE 1
Measurements of Retail Price Behavior
Five-City Statistics

Good or service	First year of series	Average annual price changes: First year to 1993	Average annual price changes: 1976 to 1993	Average coefficient of variation: First year to 1993	Average coefficient of variation: 1976 to 1993	Average annual change of CV: First year to 1993	Average annual change of CV: 1976 to 1993	Random walk process: First year to 1993	Random walk process: 1976 to 1993	"Corrected" ave. ann. chng. of CV: First year to 1993	"Corrected" ave. ann. chng. of CV: 1976 to 1993	"True test"
Alcoholic beverages	1976	-	4.82%	-	7.91%	-	+0.51%	-	Yes	-	None	No
Cereal and bakery products	1950	4.60%	5.47	5.33%	5.89	+1.49%	+2.96	Yes	?	None	+3.80%	Yes
Dairy products	1950	3.54	3.93	4.49	5.13	+0.97	-3.81	?	No	None	-3.12	No
Fruits and vegetables	1950	4.63	6.22	8.78	7.38	-1.50	+0.95	?	?	None	None	Yes
Meat, poultry, fish, eggs	1976	-	3.81	-	3.74	-	-3.98	-	?	-	None	Yes
Canned and prepared foods, baby foods, snacks, and other foods	1976	-	5.18	-	4.75	-	None	-	?	-	None	Yes
Men's and boys' clothing	1950	2.55	2.96	11.05	12.20	+0.92	-0.54	?	?	+0.30%	None	Yes
Women's and girls' clothing	1950	2.02	2.29	17.00	15.10	-0.75	-1.74	?	No	None	None	Yes
Footwear	1950	3.38	3.46	16.25	12.83	-1.50	-2.26	?	Yes	None	None	Yes
Motor fuels	1978	-	4.52	-	7.51	-	+0.76	-	?	-	None	No
Private transportation	1950	3.89	5.26	6.30	4.65	-2.30	-7.55	Yes	?	None	-5.27%	No
Piped natural gas and electricity	1950	4.27	6.08	22.38	18.13	-1.32	+0.44	?	No	None	None	No
House furnishings & operations	1976	-	3.53	-	10.33	-	+2.55	-	?	-	+2.25	Yes
Housing rents	1957	3.84	5.53	12.40	15.40	+2.18	+3.82	?	?	+2.83	None	No
Shelter (rents & rental equiv.)	1953	4.94	6.75	27.37	28.62	None	None	?	?	None	None	No
Entertainment (e.g.,books, movies, and so forth)	1976	-	4.94	-	12.25	-	+1.22	-	?	-	+1.81	No
Medical services	1967	8.26	8.56	10.85	10.71	None	+2.14	?	?	None	+1.25	No

Data sources are discussed in the text. The CV is the unweighted standard deviation. Gasoline is one part of the private transportation series; housing rents are one part of the shelter series. The meaning of "true test" is discussed in the text.

The presence of a random-walk process is determined with the use of the Dickey-Fuller test on a synthetic regression. This calculation, and the "corrected" trend calculation are described in the text.

An obvious problem arises in generalizing about price convergence or divergence. The trend values for the two time periods are quite different, showing that the trend value often depends upon the choice of the end points.

Before turning to the calculations testing for the presence of a random walk and/or persistent shocks, it is worthwhile to consider what economic factors might underlie such behavior of the coefficient of variation of prices across space. I ran a series of regressions to test whether price divergences for a given product group were related to business-cycle conditions, changes in real per capita income, changes in the CPI, and changes in the price of the commodity or service under consideration. In some cases such correlations were found, but they were not consistent with regard to sign from product group to product group. Although many economists have argued that price changes have less informational value in inflationary situations, I found no systematic evidence that spatial dispersion of prices was related to inflation.

Presence of a Random Walk Process

For the longer series from 1950 up to 1993, only 2 of the 11 series are surely random walks. For most of the other series, however, the lagged values did have statistically significant coefficients of considerable size, which means that any shock to the system tended to persist for many years. This suggests, in turn, that competitive forces and arbitrage mechanisms on the retail level impact prices oaver a number of years. For the shorter series from 1976 or beyond to 1993 (given the need to take lags into account, the first year for a number of series starting in 1976 is 1979) the statistical tests reveal either the definite presence or the absence of a random walk process in more series. Given the relatively short time period, however, not much weight can be placed on these results.

"Corrected" Trends in Spatial Variation

The most conclusive test whether long-run convergence or divergence of prices over space occurs must take the impact of the long-term shocks into account. This can be accomplished with the same regressions used to test the presence of a random walk process. For the longer series, only 2 of the 11 series show a statistically significant trend, and in both cases these prices in the five cities diverged. Focusing just on the "true test" product groups, only 1 of 5 series shows a statistically significant trend, and it indicates price divergence.

For the 1976 to 1993 period, 4 series show a statistically significant price divergence (a positive coefficient) and 2 show a significant price convergence (a negative coefficient). The remainder display no trend. For the "true test" product groups, 2 of the 8 show a significant price divergence, and the remainder have no trend.

Given the method by which these trend values are calculated, the most important feature of these calculations is the statistical significance and the sign, not the value, of the estimated trend. Although it seems possible that service prices might diverge,

as they appear to have done, other results are not so easy to interpret. It is not intuitively clear why cereal and bakery product prices diverged, while prices among dairy products converged, and other food product groups showed no significant trends in price dispersion at all in the five cities.[7]

A FINAL WORD

The strongest message from these statistical experiments is that for most product groups, the dynamic version of the law-of-one-price does not seem to hold in the post World War II period. Measurements of the variation of prices of various product groups over five cities reveal no distinct trend or actual price divergence in most cases. If market integration is measured by price variations, then such integration has not increased on retail markets in the U.S. following World War II.

These results seem attributable to two offsetting forces: on the one hand, declining transportation costs acted to decrease variation; on the other, the increasing heterogeneity of goods, a decline in the possibilities of profitable arbitrage, and the rising costs of search acted in the opposite direction. In most cases the two sets of forces appear to have nullified each other.

These empirical results are tentative because the quality of the raw data is not as high as we would want and, in addition, the statistical procedures rest on reasonable assumptions that could not, however, be verified. Much remains to be done, before this conclusion can be considered definitive. Nevertheless, I hope to have shown that the law-of-one-price may not hold, a possibility that foreign trade economists have been wrestling with for many years.

APPENDIX
A NOTE ON THE QUALITY OF CITY PRICE INDEXES

Since the city price data are so crucial to my calculations, it is worthwhile to investigate the quality of these series. Two biases are readily apparent.

Marshall Reinsdorf [1991] argues that the price chaining-process employed by the BLS when one sampled outlet is substituted for another produces a certain drift in the indexes that does not reflect actual price changes. More specifically, when a price from a branch of a discount store or outlet of a national company is substituted for the price from a small mom-and-pop store, the chaining procedure leads to an index value that does not reflect the relative decline in prices because the large stores or the outlets have lower prices. Such "outlet substitution" in the sample prices gives the BLS indexes an upward bias.

Similarly, when retail stores substitute a higher quality for a lower quality good, the ostensible price rise may not cover all of the increase in quality because the stores may wish to give customers an incentive to buy the new product by not raising prices the full extent of the quality difference. Since the BLS takes this change in quality into account only by measuring it in terms of the price increase, it introduces an upward bias into the price index.

The impact of these biases introduced by outlet or quality substitution should be greater in the indexes for particular cities, where fewer outlets and goods are sampled, than for the nation as a whole. For the long-term, however, these biases would have little impact on the price dispersion indexes discussed here if such outlet and quality substitution occurred at roughly the same rate in the different cities.

Although this seems to be a reasonable assumption, we can gain some idea of the magnitudes involved from the following consideration: Because of these biases, the price changes in the five cities for any given product group should be quite different from one month to the next. If the static law-of-one-price holds, we can determine from Monte-Carlo simulations the impact of random price differences in the five cities from the "national price" on price changes in different time periods and compare this with the price changes for the price indexes in question.

More specifically, I start with the following formula:

$$Q_t = Variance\left[\left(\frac{P_t^a}{P_{t-x}^a} - 1\right),\ldots,\left(\frac{P_t^e}{P_{t-x}^e} - 1\right)\right]$$

where P = the monthly price index for a particular city, the superscripts a, b, c, d, e designate the five cities, the subscript t designates the particular month, and the subscript x designates the lag (1 month to 48 months). These Q_t-statistics were calculated for each month for each city in the period designated by the end points; and then these calculated Qs were averaged.

Suppose, for a moment, that the static law-of-one-price held so that the price in city a at any time t would be $(N_t + e_t^a)$, where N is the "national price" and e is a random variable for city a with a mean of zero and a variance of 10 percent. Averaging the

TABLE A
Variances of Price Changes

Category/Period	First year	Last year	3 months change	6 months change	12 months change	24 months change	48 months change
Alcoholic beverages	1976	1993	0.013%	0.023%	0.044%	0.105%	0.261%
Cereal/bakery products	1950	1993	0.013	0.019	0.033	0.063	0.130
Dairy products	1950	1993	0.019	0.031	0.046	0.092	0.172
Fruits/vegetables	1950	1993	0.090	0.133	0.082	0.108	0.205
Meat,poultry, fish, eggs	1976	1993	0.018	0.027	0.039	0.077	0.136
Canned foods/misc. foods	1976	1993	0.012	0.020	0.031	0.058	0.152
Men's and boys' clothing	1950	1993	0.031	0.044	0.056	0.086	0.167
Women's and girls' clothing	1950	1993	0.086	0.122	0.135	0.212	0.351
Footwear	1950	1993	0.045	0.065	0.086	0.146	0.294
Motor fuels	1978	1993	0.071	0.109	0.114	0.146	0.179
Private transportation	1950	1993	0.014	0.021	0.032	0.055	0.096
Piped natural gas/elec.	1950	1993	0.085	0.133	0.124	0.261	0.710
House furnish/operations	1976	1993	0.013	0.018	0.026	0.041	0.079
Housing rents	1957	1993	0.004	0.007	0.021	0.072	0.259
Shelter	1953	1993	0.013	0.024	0.053	0.153	0.483
Entertainment	1976	1993	0.015	0.027	0.051	0.098	0.260
Medical services	1967	1993	0.011	0.019	0.032	0.076	0.242

The data come from the city price series described in the text. The variance calculations are described in this appendix.

results of 10 Monte-Carlo simulations, I found that for 3, 6, 12, 24, and 48-month periods, the mean price variance over a 516-month period were respectively (standardizing so that the mean price variances for 48 months = 1.00): 0.73, 0.75, 0.79, 0.84, and 1.00. For the actual price for the commodity groups used in this study, these mean price variances were (standardizing in the same way): 0.16, 0.24, 0.28, 0.49, 1.00. The results for the individual commodity groups are given in Table A. Equally dramatic differences are found if, instead of the variance of these price changes over different time periods, the coefficients of variation are calculated instead.[8]

Two explanations of the differences between the simulations and the actual price indexes can be given: either the static law-of-one-price does not hold or the series for the various cities exhibit long-term drift due to biases introduced through outlet or quality substitutions in the index. Unfortunately, sufficient research is not available to determine the degree of such drift. This essay assumes that such drift is relatively unimportant and that it is the static law-of-one-price that must be rejected. If I am incorrect, the conclusions drawn from the statistical analysis about the dynamic law-of-one-price are not correct.

NOTES

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1. I draw upon current and constant price data from NIPA that are published in *Survey of Current Business* 73, No. 6 (July 1993) and, for previous years, various volumes of U.S. Department of Commerce, Bureau of Economic Analysis, *National Income and Product Accounts of the United States*. Although the implicit GDP deflators are not strictly comparable to the CPI data (the former more closely approximate a Paasche index; the latter, a Laspeyres index) the differences in the rates of change are sufficiently great that the generalization about declining relative transportation costs seems safe. Marshall Reinsdorf of the BLS informs me that the clothing series had an unintentional downward bias in past years.
2. According to foreign trade data [International Monetary Fund, annual, various editions], the ratio of foreign trade (c.i.f. basis) to foreign trade (f.o.b. basis) fell over the period under investigation, which suggests that transportation costs as a share of foreign trade declined. National account data [U.S. Department of Commerce, 1992/93, Tables 6.1B and 1.3] show that over the same period value-added in transportation fell as a ratio of total value-added in private industry and also of the total production of all goods. These results lead to the same inference about the declining share of transportation costs in total costs of retail goods.
3. I use this particular example because it also illustrates a serious problem in calculating the hedonic indexes on which the absolute price differences are based, namely how many characteristics of a product must be taken into account to make valid absolute price comparisons. For the case at hand, how much does the New York price represent the higher percentage of designer lamps sold there than in the Connecticut suburbs of New York, rather than any absolute differences in price of exactly the same item? To minimize this problem, I have confined my sample to large metropolitan areas that presumably have populations with the same degree of sophistication.
4. It might be added that variance of the time-series has changed over time, because of not just changes in sample size but also, a change in the sampling methods in 1978. Prior to 1978 the same specific items were sampled at several outlets and the local indexes were calculated as a ratio of the average prices. After 1978 a probability method of sample selection was employed in which different items were sampled from each outlet and combined using a Laspeyres index. As a result of this 1978 change, the samples became more heterogeneous. With the regression technique I use, this should not make much difference. The qualitative results of the sample from 1950 to 1993 are not much different from those from 1976 to 1993.
5. In the statistical analysis I have assumed that the distribution of the coefficients of variation is such that standard tests of statistical significance can be employed (for instance, the *t*-test in the regression equations). Since the general interpretation of the results does not change if different significance levels are used, the errors introduced on this account should be small.

   It is also necessary to consider what bias my statistical methods might impart to the results. Since the data used in this study are average prices of aggregates, each composed of relatively substitutable products, such data should obviously reduce the spatial variation at one point in time. But it is difficult to see any major impact such aggregation should have on the *trend* of the spatial variation. More intractable is the problem arising from the weighting of the aggregates. The average prices for each aggregate are calculated for a one-year period in the late 1980s according to the CPI weights in use at that time. These were extrapolated backwards in time with the use of time series of the aggregates of current prices of the commodity group, which were periodically reweighted. It is impossible to determine the bias imparted by this procedure but to take such matters into account, I have calculated the trends not just from 1950 through 1992, but also for a shorter time period where the bias imparted by reweighting is not so strong. Finally, the results do not appear to be influenced by changes in the measurement of goods and services included. Except for rental equivalents of owner-occupied housing, major changes have not been made; moreover, any such changes would occur in all cities where the data were collected and should have less impact on the *CV* than in the price increases *per se*.
6. The regressions estimated are of the following standard form for this procedure: $X_t = a_0 + a_1 T + a_2 X_{t-1} + a_3 (X_{t-1} - X_{t-2}) + a_4 (X_{t-3} - X_{t-4})$, where the *X*s are the logarithms of the *CV*s, *T* is the year, and the lower case letters are the calculated regression coefficients. The term a_1 is the trend value; and a_2 is the coefficient on which a Dickey-Fuller test is performed to determine whether it is equal to 1 or more. The critical values at the .95 level of confidence of the statistic are reported in Fuller [1976, 373]. Technically, I am testing for the existence of unit roots. The "corrected trend" value, namely coefficient a_1, does not necessarily represent the true value of the deterministic part of the trend, but rather reflects the value of the trend when the particular type of regression is calculated; if a different specification of the synthetic regression is made, this trend value might be different. For this "corrected trend" the statistical significance and the sign are more important than the value.

   One other problem deserves mention. It is possible that the various series are influenced by a large random shock that make them appear as if the series were a random walk, when they are not. Although the calculation of two regressions with different starting points may be of some help on this problem, it is by no means satisfactory. As far as I could tell, econometricians have no satisfactory way of dealing with this problem, and neither do I.
7. Milk prices, which are governmentally controlled in a number of states, constitute only a small part of the total dairy product category.
8. For 3, 6, 12, 24, and 48-month periods, the standardized coefficients of variation in 10 different Monte-Carlo simulations for 516 months were respectively: 5.70, 7.23, 4.86, 1.48 and 1.00; for the actual series these were: 8.14, 8.12, 2.88, 1.82, and 1.00.

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