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Author(s): Jayendu Patel and Richard Zeckhauser
Source: *Journal of Business & Economic Statistics*, Vol. 8, No. 2 (Apr., 1990), pp. 179-189
Published by: Taylor & Francis, Ltd. on behalf of American Statistical Association
Stable URL: https://www.jstor.org/stable/1391980

 Journal of Business & Economic Statistics, April 1990, Vol. 8, No. 2

Shared Price Trends: Evidence From U.S. Cities and OECD Countries

Jayendu Patel and Richard Zeckhauser
John F. Kennedy School of Government, Harvard University, Cambridge, MA 02138

How similar is the price behavior of different commodities? Of the same commodity in different areas? Using cointegration methods, we explore the degree of shared trends among prices within and across U.S. metropolitan areas, as well as within and across Organization for Economic Cooperation and Development (OECD) countries. Within U.S. metropolitan areas, there are shared price trends among services, goods, and food, though not for rents. Across U.S. metropolitan areas, there appear to be empirically effective arbitraging mechanisms for the prices of food, goods, and services. Within OECD nations, there do not appear to be shared price trends between commodities, which suggests that more independent factors affect prices at the national level than at the metropolitan level. Looking at single-commodity groups across these nations, we found hardly any examples of shared trends, which suggests that substantial impediments to the free flow of goods (or factors) persist among these nations. Our results argue against the justification of a universal price index due to shared trends.

KEY WORDS: Cointegration; Consumer price index; Price aggregation.

1. INTRODUCTION

How similar is the price behavior of different commodities and of the same commodity in different areas? Although relations among trends of summary price indexes have been extensively investigated in the context of purchasing power parity, the degree of shared price trends at a more disaggregated level has been little studied. Yet an understanding of disaggregated behavior is essential to the use of summary price indexes to inform monetary policy, to determine wage increases, to protect the purchasing power of social-security recipients, or to ensure equity in tax burdens. An important question is the extent to which indexes can capture the essential long-run behavior of the components they summarize. A better understanding of the extent to which prices share long-run behavior might also shed light on the economic processes that determine price levels. Such findings, when incorporated in stylized models like that of Kydland and Prescott (1982), could help us understand the structure and unity of the economy both within an individual nation and across nations.

Previous studies on price behavior, such as those of Vining and Elterowski (1976), Parks (1978), Fama and Schwert (1979), and Gale (1981), found considerable cross-sectional heterogeneity in price *inflation*. Unfortunately, they did not inform about relations in price *levels*. Previous time series studies probably worked with inflation rather than with price levels because standard econometric test methods, such as those based on ordinary least squares (OLS), are inconsistent when the variable levels have stochastic trends, as price levels do. [See Granger and Newbold (1986) on "spurious" regressions and Nelson and Plosser (1982) and Schwert (1987) on stochastic trends in price levels.]

Traditional prefiltering to remove stochastic price trends sacrificed information on long-run behavior. In this article, drawing on recent developments in cointegration theory (see Engle and Granger 1987; Granger 1986), we provide a first look at shared trends among prices. Our definition of trend follows the cointegration literature. Briefly, any linearly regular time series characterized by a unit root can be decomposed into a random-walk component and a stationary component (Beveridge and Nelson 1981). The random-walk component is interpreted as the stochastic trend. Two series are said to share to share a trend if their trend components are proportional to each other. Equivalently, the two series have an error-correction mechanism; that is, every *permanent* shock in the trend of one series is ultimately transmitted to the trend of the other series. A shared trend implies that the error variance of the conditional forecast of a series in which the forecast is conditioned on the other series is bounded for every forecasting horizon. These concepts are readily extended to sharing of trends for more than two series.

Our study of shared trends among prices complements the literature on cost-of-living indexes (COLI's) that has extensively studied the substitution bias and heterogeneity problems [see surveys by Diewert (1983, 1987) and Pollak (1983)]. Substitution bias arises because the fixed base-period weights will not be precisely applicable as relative prices change and rational agents substitute cheaper goods for those that have become more expensive. Recent empirical studies by Braithwait (1980) and Manser and McDonald (1988) found that the substitution bias is small. The heterogeneity problem arises from differences in true costs of living among households with different demographic attributes. Kokoski (1987) reported considerable heterogeneity among demographic groups within a cohort. His results suggest that it will be difficult to produce a useful universal price index for long horizons if price trends of component bundles are not shared within an area or if price trends of a given bundle differ across geographic regions.

We explore for such within-area and cross-area empirical relations at the metropolitan level for six U.S. cities and at the national level for six Organization for Economic Cooperation and Development (OECD) countries in the 1963–1986 period. We study quarterly consumer price data for four major categories—rent (excluding imputed rent), services (less rent), goods (less food), and food (less beverages and tobacco). These categories in 1981 represented the following proportions of the U.S. consumer price index (CPI): rent = 5%, services less rent = 38%, goods less food = 40%, and food = 17%.

We address such questions as: What trends in food prices, if any, are common to Boston, Chicago, Detroit, Houston, Los Angeles, and Washington, D.C.? Shared trends, if found, might sensibly be thought of as *the* food-price trends. Similarly, if any shared trends are found in Japanese prices of rent, services, goods, and food, then they can be referred to as *the* Japanese price trends.

The article is organized as follows. Section 2 reviews some elements of index-number theory and relates it to the sharing of price trends. The relevant cointegration tests are discussed. Section 3 applies the tests to the component prices within and across six U.S. metropolitan areas. Section 4 extends the empirical examination to prices at the national level within and across six OECD countries. (Exchange rates are admitted as explanatory variables to avoid imposing long-run purchasing-power parity.) Section 5 presents conclusions.

2. SHARED PRICE TRENDS: MOTIVATION, THEORY, AND TEST METHODS

Knowledge of shared price trends provides insights into the structure of the economy. For instance, King, Plosser, Stock, and Watson (1987) sketched a general equilibrium model in which the concept of shared trends is central. The notion of shared trends can also be important in the context of COLI's, especially since, as Deaton and Muellbauer (1980), Michael (1979), and Kokoski (1987) emphasized, the differences in states and natures of consumers lead to pitfalls in naive construction of a universal COLI. Nonetheless, undifferentiated price indexes like the CPI are in widespread use, vitally affecting the welfare of millions by, for example, determining increases in wages and social-security payments.

Are we then guilty of using indexes inappropriately? Have we failed to heed Von Mises (1933) who pointed out that "He who cares to go to the trouble of demonstrating the uselessness of index numbers . . . will be able to select a good proportion of his weapons from the writings of the very men who invented them" (p. 188)? Besides the economic criteria advanced in the literature for assessing indexes, another can be based on the statistical property of shared trends; we show that a universal price index is beneficial if price trends are shared and if a market for trading on a standard price index exists.

2.1 Shared Price Trends and Cost-of-Living Indexes

Consider the following stylized situation. The COLI for agent A is C_i^A, where state i has occurred. For example, i may be a state in which the price of oil is high. The corresponding COLI for agent B is C_i^B. If $C_i^A = \alpha C_i^B$, where α does not depend on the state that actually occurs, then the existence of a market that permits trades contingent on the outcomes of C_i^A will be complete in the sense of enabling agent B to hedge his or her cost of living. Now suppose that there are more than two agents. Without loss of generality, we select the reference agent to be 1. If the COLI for any agent k is simply $\alpha_k C^1$, then C^1 can be a representative COLI for all agents. [Fama and Schwert (1979) advanced a scenario that implies constant relative logarithmic prices for components; it also admits a meaningful representative price index.] A detailed model is now sketched.

Denote the income–price state of a household, h, by $(Y, \mathbf{P})$, where Y is the household's income and $\mathbf{P}$ is the price vector of commodities. Following Theil (1980), for instance, let us evaluate the consequence of an infinitesimal change $(dY, d\mathbf{P})$. A cost-of-living comparison at the prevailing indirect utility level before the change, $U_l(Y, \mathbf{P})$, can be written as

$$\log C^h(\mathbf{P} + d\mathbf{P}, \mathbf{P} \mid U_l(Y, \mathbf{P})) = \sum_{i=1}^{N} w_i d(\log P_i), \tag{1}$$

where C^h is household h's COLI, w_i is the weight for commodity i in the household's consumption basket, and P_i is the price of commodity i; the summation is over i, where i runs over the N goods relevant to the consumer. Expression (1) is simply the Divisia price index. In applications, the differential approach is approximated by the discrete version. The standard situation considers a COLI linking this period to the previous period.

Denote by Δ the backward difference operator (i.e., for any variable x_t, $\Delta x_t \equiv x_t - x_{t-1}$). The discrete Divisia index due to Tornqvist, setting all lower-case price variables to denote log transforms of the corresponding price levels, is

$$\Delta c_t^h = \sum_{i=1}^{N} \omega_{it}^h \Delta p_{it}. \tag{2}$$

Here ω_{it}^h is the weight that applies to commodity i and is defined as $[(P_{it}Q_{it}^h/\ \mathbf{P}_tQ_t^h) + (P_{i,t-1}Q_{i,t-1}^h/\ \mathbf{P}_{t-1}Q_{t-1}^h)]/2$, Q^h is the quantity vector of purchases made by the "rational" household, and $\mathbf{P}$ is the price vector facing the consumer. Although the Tornqvist discrete index is exact only for a translog utility function, Diewert (1978) showed that it provides a second-order approximation to an arbitrary twice-continuously-differentiable aggregator (utility) function subject to a budget constraint.

Given an index like Equation (2) for the household, it is still difficult to define a representative index because ω_{it}^h differs across households. Invoke the severe simplifying assumption that the ω_{it}^h is time-invariant; that is, $\omega_{it}^h = \omega_i^h$ for all t. The small substitution biases measured in the literature indicate that this may not be a bad approximation. Set $t = 0$ as a reference time period, and hereafter consider all prices to be relative to this reference date. Suppose that the log-transformed prices of individual goods, p_{it}, are individually integrated of order 1; that is, a first-order difference is needed to render them stationary. This assumption is empirically reasonable (see Sec. 3). Consider a condition that the $p_{.t}$ are pairwise cointegrated; that is, there exists an α_{ij} such that $e_{ijt} \equiv p_{it} - \alpha_{ij}p_{jt}$ is stationary for all i and j though the individual $p_{.t}$ are mean nonstationary. Note that the stationary restriction on e_{it}, $I(0)$ behavior, is strong; generally the order of integration of e_{it} will be the same as that of $p_{.t}$—that is, $I(1)$. For instance, such cointegration could occur if permanent shocks to prices are of monetary origin and the quantity theory of money holds with a log-linear demand-for-money schedule. Typically, the cointegration condition rules out permanent productivity shocks that are *idiosyncratic* to a component.

With commodity 1 as the numeraire, the assumption of cointegrated prices implies

$$p_{it} = \alpha_i p_{1t} + e_{it} \quad \text{for all } i, \tag{3}$$

where e_{it} is stationary. Combining the assumptions and the normalization in Equation (3), we obtain

$$p_t^h = p_{1t} \sum_{i=1}^{N} \omega_i^h \alpha_i + \sum_{i=1}^{N} \omega_i^h e_{it}. \tag{4}$$

Equation (4) states that the logarithms of the COLI's of different households are cointegrated or, equivalently, that they share a trend.

The fortuitous conditions needed for such a representative COLI to apply are unlikely to arise in practice, but similar conditions may be approximately met. If relative prices of relevant commodities are in a suitable long-run relation, then policies or financial instruments based on a representative COLI will be useful for agents who have a long-time horizon. For instance, if the path of the quantity of money is the sole determinant of trends in nominal prices and if the path affects all prices similarly, all prices will move in lock-step over the long run. (*Short-run* deviations may arise because a shift in the quantity of money will probably be transmitted to some goods' prices more quickly than to others.) Alternatively, if price trends are driven by the technological evolution in a basic factor of production, price trends (after removing deterministic components and making a suitable transformation) may be proportional to each other in the long run.

The following example sketches one practical implication arising from Condition (4). Agent B wishes to hedge cost-of-living uncertainty, say because multiperiod wage contracts are fixed in nominal terms or are tied to a standard-price index that is not the same as that of the agent. Suppose the only available futures contract on price indexes is based on the true COLI of agent A. Obviously, this enables A's price risk to be perfectly hedged. Agent B, however, differs from A in that the weights of items in their consumption baskets are different. Under the assumption of cointegrated relative prices, the hedging effectiveness B can attain, conventionally measured by the percentage of variance reduction achievable (see Ederington 1979), approaches 100% as the horizon over which the hedging performance is measured increases. In other words, for a horizon τ, the inferiority of the hedge per unit time available to B relative to A goes to 0 as fast as τ goes to infinity. Favorable evidence for short hedging periods (90 days) is reported in related work on hedging COLI risk (where COLI is measured by CPI) using existing treasury-bill futures contracts (Patel and Zeckhauser 1989).

More generally, agents' preferences will reasonably be such that $\omega_{it}^h \neq \omega_{ik}^h$ for some t and k. In this case, shared price trends across components will suffice to hedge individuals' inflation risks only if a complete set of options contracts on the fundamental price trend is available.

The discussion—so far framed in terms of price trends of components—readily extends to the case of price trends of a given consumption bundle across regions. The index subscript i in Equation (3) can be interpreted to apply to N different regions. We now turn to the problem of testing whether observed prices indicate a sharing of trends as expressed in Equation (3).

2.2 The Cointegration Test Methods

Our maintained hypothesis is that all price series we deal with, after suitable log-transformation, are linearly regular processes. For such series, the hypothesis of univariate stochastic trends can be tested following Dickey and Fuller (1981). The null hypothesis for their test [generally called the augmented Dickey–Fuller (ADF) test], is that the price series, p_t, has a unit autoregressive root (i.e., has a stochastic trend). The ADF test is based on the t statistic associated with the ρ coefficient in the following regression estimated by OLS:

$$\Delta p_t = \rho p_{t-1} + \sum_{i=1}^{l} \beta_i \Delta p_{t-1} + \varepsilon_t. \tag{5}$$

Here l is selected to be large enough to ensure that ε_t is a white-noise series. The null hypothesis of stochastically trending p_t is rejected if ρ is negative and significantly different from 0. The distribution of the t test for ρ is *not* the usual Student-t distribution under the null hypothesis; rather, it is that given by Fuller (1976).

In practice, we rarely know l. Said and Dickey (1984) showed that the ADF test is valid asymptotically if l is increased with sample size (T) at a controlled rate ($T^{1/3}$). For our sample size of $T = 100$, this translates into $l = 4$. Since setting l unnecessarily large reduces the power of the tests, however, we select l by a simple pretest. The null hypothesis of our pretest is that l is 0; the alternative hypothesis is that l is 4 (which is adequate in all our cases to render ε_t an empirical white noise). The critical value for the pretest was a 10% significance level. An alternative approach would be to use a model-selection procedure based on some information criterion (Engle and Yoo 1987, p. 157). Schwert (1987) pointed out also that when the series has a moving average component the ADF test can perform inconsistently in small samples unless l is large, which would reduce the power of the test. Given these limitations of the ADF test, we always buttress it with a qualitative examination of the autocorrelogram of p_t—mean stationarity is indicated if the autocorrelations decay rapidly at higher lags; otherwise, a trend is present.

Compared with the ADF test, a Durbin–Watson-based test has more power. Its validity, however, is restricted to cases in which the alternative hypothesis is a stationary (autoregressive) AR(1) process. Such an alternative is inappropriate given the empirical dynamics of prices. Hence we do not employ Durbin–Watson–based tests.

Once a stochastic trend is confirmed for the individual price series, we explore for shared trends among series using methods recommended by Engle and Granger (1987). Consider the null hypothesis that there are no shared trends between two series, $p_{1,t}$ and $p_{2,t}$. The alternative hypothesis is that they share a common trend with the proportionality between the two trends equal to γ. Under the null hypothesis, the series $z_t \equiv p_{1,t} - \gamma p_{2,t}$ will be trending, whereas z_t will be stationary under the alternative. In this case, the ADF test applied to z_t is appropriate.

More generally, there are N price series, denoted by $p_{1,t}, p_{2,t}, \ldots, p_{N,t}$, each of which has a stochastic trend. These trends may be shared (i.e., they may not be distinct). We say that n series *share m trends iff the trends have some commonality—that is, if m is less than n;* if m equals n, we say that there are no common trends. [Note that our terminology differs from that of Engle and Granger (1987), who would say that there are n common trends.]

Let $p_{1,t}$ be the reference series. If the trend of $p_{1,t}$ is linearly related to the other $N - 1$ trends, then there exists a unique coefficient vector, β, such that $z_t \equiv p_{1,t} - \sum \beta_i p_{i,t}$ is a stationary series. Of course, β may not be known a priori. Stock (1987) showed that β can be consistently estimated using OLS in the following regression:

$$p_{1,t} = \sum_{i=1}^{N} \beta_i p_{i,t} + z_t. \tag{6}$$

If $p_{i,t}$ does not share a trend with the other $N - 1$ prices, then the OLS estimate of β is "spurious" and z_t is nonstationary. A test for the null hypothesis that there is no shared trend (i.e., there are N distinct trends) can be based on testing for nonstationarity of the regression residuals [z_t in Eq. (6)]. The ADF test can be employed. We refer to the test from the two-step approach [of first estimating Eq. (6) and then applying the ADF test to the regression residuals] as the ADF2 test. The appropriate limiting distribution of the t statistic is no longer that of the ADF test, since the ADF2 test is based on regression residuals. Asymptotic percentile values for the ADF2 test, which depend on the included variables in the first-stage regression (6), can be inferred from the results reported by Engle and Yoo (1987, table 2). It is tempting to use their critical values that take into account additionally both sample size and small-sample dependence of the ADF2 test on the number of estimated parameters in the second-stage regression (Engle and Yoo 1987, table 3). We find it unreasonable to maintain the hypothesis that the price series we examine are generated by the specific models used in their simulations, however.

To establish the minimum number of trends characterizing a group of prices, the ADF2 tests, reported in Sections 3 and 4, have to be applied sequentially to decreasing subsets of prices. In contrast, Stock and Watson (1987) offered a direct method to determine the minimum number of trends. Their approach, however, yields little intuition about the underlying trends. Generally, the application of the Stock and Watson methods to the price series in this article gives results (not reported) similar to those using the ADF2 test. For instance, the two methods identify the same number of shared trends for each cross-city relation. When the ADF2 approach finds no shared trends, however, there are some differences—the Stock–Watson approach indicates one or two fewer trends, almost never indicating no shared trends. An evaluation of the few instances in which the test methods yield different results is beyond the scope of this article.

3. PRICE BEHAVIOR WITHIN AND ACROSS U.S. METROPOLITAN AREAS

For the post-1967 period, we examine the price behavior of rent, services, goods, and food in the metropolitan areas of Boston, Chicago, Detroit, Houston, Los Angeles, and Washington, D.C. The data, not seasonally adjusted, were compiled by the U.S. Bureau of Labor Statistics. All series are transformed to logarithms and prefiltered by regression on a time trend and a constant. The prefiltered series are consistent with mean nonstationarity, as in other reports on broad indexes in the literature—for instance, Nelson and Plosser (1982) or Harvey (1985). In results not shown, the autocorrelations decay very slowly, and ADF tests fail to reject mean nonstationarity.

Table 1 shows the autocorrelations of the inflation rates. Generally, the inflation rates appear to be stationary, with the autocorrelations close to 0 by lag 8 (see the Boston price series, for instance). The possible exceptions among the 24 inflation series appear to be Houston, rent; Houston, services; Los Angeles, rent; Los Angeles, goods; and Washington, goods. The inferences based on the ADF results are mixed. Once we allow for the strong seasonal component in Washington rent and services, however, and weigh the information from the autocorrelogram, only the five series identified as exceptions appear to exhibit a possibly small trend. We conclude that the price series, with rent as a possible exception, are reasonably characterized as stationary in the inflation rates.

3.1 Shared Price Trends Within U.S. Metropolitan Areas

We search for convenient shared trends among price groups within individual metropolitan areas, following our discussion in Section 2. The ideal situation would have all price pairs in a relation like Equation (3), implying that there is but one shared trend. But the univariate results in Table 1 already foreshadow negative findings. If rent is $I(2)$ with the other prices as $I(1)$, then Equation (3) cannot hold between rent and the other prices.

The simplest case arises if, within each metropolitan area, all component prices have the same trend with unit proportionality [i.e., $\alpha = 1$ in Eq. (3)]. The ADF tests for this proposition apply to $p_{j,t} - p_{k,t}$ for all j and k within each metropolitan area. The test results are given in Table 2. Generally, the ADF tests do not reject the absence of shared trends with unit proportionality. The exception is Houston, where the alternative hypothesis is accepted.

The results of Table 2, however, may indicate only that the factor of proportionality for shared trends among metropolitan area prices is generally different from unity. Moreover, the simple approach of Table 2 cannot assess situations in which the individual price trends are a composite of more than one fundamental trend.

Consider the regression of the rent price index on the other three price indexes. Under the null hypothesis that the rent trend is not linearly related to the other price trends in the city, the residuals of the regression will be nonstationary, a matter to be examined by the ADF2 test. These ADF2 results appear in Table 3. For

Table 1. Autocorrelations and ADF Tests for Inflation Rates: Six U.S. Metropolitan Areas, 1967:1–1987:3

Inflation series	Autocorrelations at lags								ADF test*
	1	2	3	4	5	6	7	8	
Boston									
Rent	.27	.14	.16	.06	.13	.02	.16	.29	−2.71
Service	.15	.13	.03	.28	.06	.07	.15	.04	−2.61
Goods	.20	.27	.11	.24	.05	.27	.01	.03	−2.72
Food	.09	.14	.09	.36	.02	.05	.11	.04	−2.73
Chicago									
Rent	.01	.03	.21	.25	.05	.17	.01	.14	−2.69
Service	.17	.20	.09	.56	.10	.16	.13	.29	−1.99
Goods	.14	.25	.09	.30	.08	.10	.06	.18	−2.48
Food	.21	.21	.25	.30	.01	.06	.13	.03	−2.69
Detroit									
Rent	.16	.54	.09	.31	.00	.33	.09	.11	−2.45
Service	.22	.31	.37	.21	.30	.13	.01	.09	−2.00
Goods	.23	.14	.15	.20	.13	.05	.01	.07	−2.56
Food	.19	.13	.16	.33	.13	.02	.08	.12	−3.26
Houston									
Rent	.03	.49	.26	.28	.20	.08	.14	.03	−1.62
Service	.27	.32	.25	.55	.23	.11	.10	.32	−1.25
Goods	.16	.41	.13	.29	.03	.17	.02	.08	−2.65
Food	.16	.35	.18	.33	.10	.17	.14	.01	−2.95
Los Angeles									
Rent	.10	.41	.17	.59	.19	.21	.14	.31	−1.74
Service	.31	.21	.22	.13	.02	.09	.21	.02	−3.13
Goods	.16	.28	.26	.24	.15	.03	.10	.04	−2.24
Food	.23	.26	.30	.40	.02	.07	.15	.06	−2.49
Washington									
Rent	.33	.51	.21	.48	.12	.31	.16	.40	−2.49
Service	.12	.47	.16	.46	.23	.42	.22	.27	−2.51
Goods	.15	.41	.11	.31	.08	.19	.02	.07	−2.35
Food	.20	.02	.17	.22	.04	.03	.16	.16	−2.82

* See discussion of the test in Section 2.2. The 5% (10%) critical value of the ADF test under mean nonstationarity is −2.9 (−2.6).

Table 2. ADF Test for Unit-Proportional Price Trends Within Six U.S. Metropolitan Areas, 1967:1–1987:3

Inflation series	Service	Goods	Food
Boston			
Rent	−.99	−1.42	−1.78
Service		− .97	−2.48
Goods			−2.14
Chicago			
Rent	−2.27	− .26	−1.23
Service		−1.27	−1.80
Goods			−1.50
Detroit			
Rent	−1.37	−1.84	−1.89
Service		−2.75*	−1.50
Goods			−2.59*
Houston			
Rent	−3.39*	−4.87*	−3.61*
Service		−3.58*	−2.73*
Goods			−2.05
Los Angeles			
Rent	−.95	−1.09	−2.12
Service		−1.53	−1.25
Goods			−3.72*
Washington, D.C.			
Rent	−.99	−1.38	−1.52
Service		−2.16	−1.33
Goods			−1.71

NOTE: The 5% (10%) critical value of the ADF test is −2.9 (−2.6). The null hypothesis is the *absence* of unit-proportional stochastic trends among prices.
* Test values that reject the null hypothesis.

every metropolitan area, we fail to reject (at a 10% significance level) the null hypothesis of nonstationary residuals; that is, we cannot reject the null hypothesis of no shared trends betweeen rent and the other three price indexes. The Houston results are closest to a finding of shared trends, which appears consistent with Table 2.

The results in Table 3 do not preclude a shared trend between prices among the subset of services, goods, and food. Table 4 reports the ADF2 tests that are based on the regressions of the services-price index on prices of goods and food for each metropolitan area. In contrast to Tables 2 and 3, the ADF2 test now *rejects* the null hypothesis of *no* shared trends. The evidence favors at most two trends between the three price series for five of the six metropolitan areas; Los Angeles is the exception.

Is there one shared trend or are there two shared trends for these three components in the six different areas? We employ the ADF2 test with bivariate regressions between prices. In this case, the null hypothesis is the presence of two distinct trends. Now Los Angeles is the only area where one trend suffices for goods and food. (See the lower panel of Table 4.) Thus the combined bivariate and trivariate results reveal a pattern that applies to each of the six areas—there are but two price trends among services, goods, and food; the rent-price index never shares a trend with the other prices. These findings are consistent with a variety of economic models. For example, the price trends of two factors, say labor and energy, determine price trends for services, goods, and food within a region, but price trends in rents reflect evolving scarcity values for location.

3.2 Shared Price Trends Across U.S. Metropolitan Areas

In Section 3.1, we found limited prospects for combining price indexes within metropolitan areas; three trends are required to preserve trend information on the four components. If price indexes *across* metropolitan areas share trends, we could usefully combine area-specific price indexes to form multiple-city or national price indexes. We turn to cross-area relations.

Table 5 reports a simple statistical evaluation based on the ADF test. For food, all bivariate pairs with Washington reject the absence of a shared trend. Of course, this implies, in turn, that all other bivariate pairs must also share a trend. The failure to observe this result in the statistical tests may be for two reasons. First, the power of the ADF test may be insufficient to obtain consistent rejections for all pairs. Second, the assumption of unit proportionality of trends may be best approximated with the Washington trend as the reference trend. The latter explanation is quite plausible. For prices of services, goods, and rent, no simple relation is suggested. In any case a direct relaxation of unit proportionality between trends is examined in Table 6.

Confirming the impression from Table 5, food prices strongly indicate a single underlying fundamental trend.

Table 3. Rent Does Not Share Trend With Services, Goods, and Food: Six U.S. Metropolitan Areas, 1967:1–1987:3

Metro area	Autocorrelations of cointegration regression residuals at lags 1	2	3	4	8	ADF2 test Value[a]	l[b]
Boston	.80	.64	.52	.37	−.15	−3.27	0
Chicago	.87	.74	.63	.52	.00	−2.95	4
Detroit	.72	.61	.54	.34	−.18	−2.31	4
Houston	.75	.66	.49	.40	−.19	−3.51	0
Los Angeles	.84	.65	.48	.34	−.09	−2.96	0
Washington	.71	.67	.43	.46	.14	−2.60	4

[a] The one-sided 5% (10%) critical value for the ADF2 test is −4.2 (−3.9) for a four-variable set. The null hypothesis of no cointegration between rent and the other price indexes is not rejected. (The ADF2 test is discussed in Sec. 2.2.)
[b] The autoregressive-correction order following Equation (5) is denoted by l.

Table 4. *Shared Trends Among Services, Goods, and Food: Six U.S. Metropolitan Areas, 1967:1–1987:3*

Metro area	β_1	β_2	Residual autocorrelations at lags					ADF2 test	
			1	2	3	4	8	Value[a]	l[b]
Services on goods and food									
Boston	.50	−.38	.66	.52	.33	.21	−.22	−4.33	0
Chicago	1.10	−.42	.57	.43	.14	.15	−.12	−3.52	4
Detroit	1.04	−.21	.74	.54	.35	.18	−.12	−3.52	0
Houston	1.13	−.06	.77	.54	.33	.19	−.37	−4.32	4
Los Angeles	1.25	−.57	.79	.67	.48	.35	−.01	−3.23	0
Washington	.87	−.25	.57	.43	.13	.06	−.20	−4.44	4
Goods on food (the only case of cointegration not rejected)									
Los Angeles	.86	not applicable	.88	.78	.63	.49	.00	−3.56	0

NOTE: There is no case of cointegration between the services price index and the food price index.
[a] The one-sided 5% (10%) critical values for the ADF2 test are −3.37 (−3.02) for a two-variable set and −3.78 (−3.47) for a three-variable set. (The ADF2 test is discussed in Sec. 2.2.)
[b] The autoregressive-correction order following Equation (5) is denoted by l.

In Table 6, for food, the null hypothesis being tested by the ADF2 test is that there are two distinct trends between Washington and each of the other metropolitan areas (i.e., no shared trend). The ADF2 test rejects (5% significance level) the null hypothesis for Boston, Chicago, and Detroit; they share their trend with Washington (and hence with each other). For Houston and Los Angeles, we can reject the null hypothesis at the 10% significance level, though not at a 5% significance level. The assertion that Houston and Los Angeles do not share a single trend can be rejected at the 5% significance level, however. Thus food prices across the six metropolitan areas are confirmed to share a single trend.

Table 5. *ADF Test for Unit-Proportional Price Trends Across Six U.S. Metropolitan Areas, 1967:1–1987:3*

	Boston	Chicago	Detroit	Houston	Los Angeles
Rent					
Chicago	−2.31				
Detroit	−2.15	−2.35			
Houston	−1.75	−2.06	−3.09*		
Los Angeles	−1.95	−1.92	−1.51	−1.38	
Washington	−2.07	−2.63*	−2.31	−1.71	−1.58
Services					
Chicago	−2.15				
Detroit	−2.18	−1.09			
Houston	−1.38	−.55	−2.21		
Los Angeles	−2.22	−2.53	−2.78*	−.20	
Washington	−2.72*	−1.66	−2.37	−1.38	−2.41
Goods					
Chicago	−2.44				
Detroit	−5.03*	−1.33			
Houston	−2.46	−1.05	−2.31		
Los Angeles	−3.54*	−2.17	−3.45*	−2.45	
Washington	−2.07	−2.46	−4.43*	−2.30	−3.50
Food					
Chicago	−2.05				
Detroit	−2.32	−4.15*			
Houston	−1.81	−2.37	−3.24*		
Los Angeles	−1.90	−1.97	−2.49	−2.80*	
Washington	−2.97*	−4.15*	−2.67*	−3.36*	−3.04*

NOTE: The 5% (10%) critical value of the ADF test is −2.9 (−2.6). The null hypothesis is the *absence* of unit-proportional stochastic trends among prices.
* Test values that reject the null hypothesis.

Table 6 does not alter the unfavorable indications for rent obtained from Table 5. Rent prices support the notion of an individualistic price trend for each metropolitan area, presumably because local conditions are a major influence on rent levels. For instance, we cannot reject the hypothesis that Boston's rent trend is distinct from a composite of the rent trends of Chicago, Detroit, Houston, Washington, and Los Angeles at 10% significance levels. Similar conclusions apply to the other cities.

Consider service prices next. Table 6 reports on the null hypothesis of three distinct trends between Chicago and Houston and each of the other metropolitan areas taken one at a time. This null hypothesis is rejected at a 10% significance level for Boston, Detroit, and Washington and only marginally not rejected for Los Angeles. In results not reported, we find that no bivariate pair rejects the null hypothesis of two distinct trends. Thus service prices across the six metropolitan areas appear to be captured by two trends.

Results for goods-price indexes are similar to those for services. Table 6 reports on the null hypothesis of three distinct trends between Chicago and Houston and each of the other metropolitan areas. The null hypothesis is rejected (at 5% significance levels) for each metropolitan area. Goods prices across the six metropolitan areas can be characterized by two trends. Two logical extensions of these findings are not taken up in this article. First, do these results apply when additional metropolitan areas are examined? Second, what are the economic causes for the numbers of trends actually observed?

Note that the finding of two shared trends for services and goods readily explains an absence of simple bivariate relations (see Table 5), which presumes one shared trend. Similarly, the finding of one shared trend with a cointegrating coefficient frequently close to unity for

Table 6. Shared Price Trends Across Six U.S. Metropolitan Areas, 1967:1–1987:3

Left side variable	Right side variables (coefficient)	Residual autocorrelations at lags 1	2	3	4	8	ADF2 test Value[a]	l[b]
			Rent					
Boston	C, D, H, L, W	.70	.69	.55	.41	−.08	−2.99	4
Chicago	D, H, L, W	.34	.49	.18	.30	−.16	−3.24	4
Detroit	H, L, W	.75	.69	.61	.51	−.08	−2.88	4
Houston	L, W	.91	.85	.75	.64	.23	−1.97	4
Los Angeles	W	.72	.72	.62	.59	.22	−2.19	4
			Services					
Boston[c]	C (.48), H (.01)	.59	.48	.30	.30	−.34	−3.74	4
Detroit[c]	C (.56), H (.39)	.68	.55	.42	.32	−.25	−3.72	4
Los Angeles	C (.73), H (.21)	.74	.55	.35	.25	−.14	−3.47	0
Washington	C (.43), H (.20)	.56	.35	.08	.07	−.12	−4.10	4
			Goods less food					
Boston								
Detroit	C (.49), H (.46)	.60	.48	.27	.21	.09	−4.25	0
Los Angeles	C (.00), H (.88)	.42	.41	.15	.15	−.17	−5.72	0
Washington	C (.40), H (.40)	.32	.08	−.09	−.06	.03	−6.43	0
			Food					
Boston	W (.84)	.73	.54	.39	.28	−.10	−3.55	0
Chicago	W (1.06)	.58	.35	.20	.17	−.20	−4.66	0
Detroit	W (1.02)	.66	.40	.23	.28	−.09	−4.04	0
Houston	W (.99)	.78	.64	.51	.48	.11	−3.34	0
Los Angeles	W (.89)	.83	.71	.62	.59	.27	−3.10	0
Los Angeles	H (.89)	.54	.53	.22	.24	−.07	−3.70	4

NOTE: Each metropolitan area is denoted by the first letter of its name in the second column.

[a] The one-side 5% (10%) critical values for the ADF2 test are −3.37 (−3.02) for a two-variable set, −3.78 (−3.47) for a three-variable set, and below −4.2 (−3.9) for more than three variables. (The ADF2 test is discussed in Sec. 2.2.)

[b] The autoregressive-correction order following Equation (5) is denoted by l.

[c] The autocorrelations suggest strong seasonality. The cointegration results are not materially different if a seasonal prefilter is applied.

food-price indexes is consistent with the findings in Table 5. Though the ADF2 approach could have substantially less power than the ADF approach when the bivariate unit proportionality condition holds, we are reassured to find that conclusions from the two test approaches are not materially different.

In sum, the evidence indicates that the markets for foods, goods, and services are sufficiently integrated within the United States that shared trends are observed across quite heterogeneous and distant metropolitan areas. As might be expected, given the immobility of land and physical structures, rental costs do not share trends across areas. There is limited evidence that price series across components share trends within metropolitan areas.

4. PRICE BEHAVIOR IN AND ACROSS OECD COUNTRIES

We now extrapolate to the level of national price indexes, drawing data from the United States, Canada, West Germany, Japan, the Netherlands, and the United Kingdom. As in Section 3, the raw data are seasonally unadjusted price indexes of rent, services, goods, and food. Figures were drawn from various issues of the OECD *Main Economic Indicators* for the post-1963 period.

We are interested in two basic questions. First, are the price trends shared within each nation? A positive finding would readily justify the widespread construction and use of broad national price indexes like the CPI and provide suggestive data for hypotheses about the structures of these national economies. Second, are there shared trends for the same subaggregates across countries? If so, this might indicate that trade is relatively free, or at least that the factors impeding trade have been relatively constant in their price effects over time.

All series are log-transformed and prefiltered by regression on a time trend and a constant. In addition, several series of Japan, the Netherlands, and the United Kingdom were observed to have a strong seasonal component. For such series, we apply a mechanical deseasonalization filter that follows Sims (1974). Briefly, the log-level series was first detrended. Next the detrended series was transformed into the frequency domain. The seasonal frequency was "masked" (i.e., zeroed). Thereafter, this masked series was transformed back into the time domain. The deseasonalized series are identified in Table 7.

The sample autocorrelations of the detrended logarithms of consumer prices for the six countries (not reported) decay very slowly, which confirms mean nonstationarity. Table 7 reports the autocorrelations of the inflation rates. Unlike the parallel Table 1, Table 7 shows substantial persistence in inflation rates. Canada and the United States have only *one* subaggregate that

Table 7. Autocorrelations and ADF Tests for Inflation Rates: Six OECD Nations, 1963:1–1986:2

	Autocorrelations at lag								ADF
	1	2	3	4	5	6	7	8	test[a]
United States									
Rent	.86	.77	.80	.82	.72	.65	.66	.66	−1.57
Services	.68	.54	.53	.45	.38	.31	.20	.18	−2.30
Goods	.45	.46	.34	.38	.21	.23	.14	.10	−1.66
Food	.37	.34	.28	.37	.02	.14	.09	.13	−3.20
Canada									
Rent	.49	.31	.32	.63	.27	.08	.04	.25	−2.30
Services	.45	.33	.38	.46	.26	.11	.14	.16	−2.29
Goods	.63	.71	.58	.65	.51	.45	.37	.35	−1.69
Food	.06	.11	.10	.46	−.06	.02	−.03	.33	−2.67
Germany									
Rent	.38	.02	.17	.35	.09	−.04	.14	.36	−2.71
Services	−.05	.02	.05	.06	.04	.03	.10	.10	−3.50
Goods	.46	.17	.33	.51	.24	.02	.16	.24	−1.38
Food	−.14	−.21	−.07	.61	−.08	−.21	−.18	.58	−2.31
Japan									
Rent	.59	.43	.48	.31	.22	.17	.14	.20	−2.32
Services[b]	.35	.38	.26	.31	.22	.05	.20	.15	−2.23
Goods[b]	.48	.39	.32	.26	.04	.00	−.02	.13	−3.27
Food[b]	.08	.08	.27	.16	−.04	.08	.22	−.15	−3.04
Netherlands									
Rent[b]	−.35	.01	−.12	.13	−.15	.23	−.01	.03	−6.12
Services[b]	.13	.19	.15	.38	.07	.20	.12	.25	−2.01
Goods[b]	.30	.23	.20	.11	.04	.15	.16	.21	−2.82
Food	−.06	−.10	−.04	.21	.09	−.02	.05	−.05	−3.16
United Kingdom									
Rent[b]	.27	.11	.09	.10	.20	.16	.14	.00	−2.82
Services[b]	.47	.27	.15	.33	.10	−.01	−.07	.22	−3.12
Goods					Not available				
Food	.34	.15	.33	.44	.25	.17	.28	.46	−1.97

[a] See discussion of the test in Section 2.2. The 5% (10%) critical value of the ADF test under mean nonstationarity is −2.9 (−2.6).
[b] These series are seasonally adjusted using Sims's (1974) method.

appears stationary (though that one, food, is the same for both countries). Generally, the other countries have two or three subaggregates that are stationary in inflation rates. These results make it unlikely that there is a single shared trend within a nation (since a necessary condition is that the integration orders of the subaggregate trends be identical).

4.1 Shared Price Trends Within OECD Nations

Results based on ADF tests (see Tables 2 and 5) are not shown here; they find no evidence of bivariate shared trends with unit proportionality. ADF2 tests on the number of shared trends are presented in Table 8. The approach follows that applied in Section 3. The null hypothesis—that there are no shared trends within each country—is generally not rejected even at a 10% significance level. The only exception is between prices of services, goods, and food in Canada, which seems to be explained by two shared trends.

These results indicate that component price trends within a nation are not principally determined by just one or two factors, such as the price of labor or the quantity of money. Our empirical findings suggest, following the discussion in Section 2.1, that a single index may not provide suitable cost-of-living hedges for heterogeneous agents in the regions studied. This may explain the lackluster reception of the CPI futures contract introduced recently in the United States. But there might be a role, at least theoretically, for futures contracts on disaggregated price indexes that let hedgers mix and match contracts to fit their circumstances.

Further, the findings suggest that it would be difficult to construct enforceable long-term wage contracts based on a broad price index. The absence of a strong relation among relative prices suggests that an index of the value of the marginal product of labor (VMPL) in a specific industry, A, is likely to drift apart from the VMPL in another industry, B. If wage-escalation clauses are in-

Table 8. ADF2 Tests for Shared Price Trends Within Six OECD Nations, 1963:1–1986:2

Nations	Four-variable set: rent, services, goods, food	Three-variable set: services, goods, food	Two-variable set: goods, food
United States	−1.66 (4)[a]	−3.02 (4)	−2.43 (4)
Canada	−2.43 (4)	−3.97* (4)	−2.39 (4)
Germany	−2.58	−2.53	−3.52* (4)
Japan	−2.79	−2.12	−2.33
Netherlands	−7.39*	−3.44	−2.80
United Kingdom[b]	—	−2.11	−3.03

NOTE: The one-sided 5% (10%) critcal values for the ADF2 test are −3.37 (−3.02) for a two-variable set, −3.78 (−3.47) for a three-variable set, and below −4.2 (−3.9) for more than three variables. The rejections of the null hypothesis are marked by asterisks in the table. (The ADF2 test is discussed in Sec. 2.2.) The dash indicates "not available."
[a] The numbers in parentheses refer to a nonzero l in the autoregressive-correction order following Equation (5).
[b] Data on food prices were unavailable for the United Kingdom only. For the United Kingdom, the three-variable set is rent, services, and goods, and the two-variable set is services and goods.

dexed to a common COLI, wages will not track this divergence. Efficiency will be lost. This difficulty may partially explain the absence of long-term wage contracts in practice.

4.2 Shared Price Trends Across OECD Nations

We next review some findings on the relation among the prices for the four components—rent, services, goods, and food—*across* nations. As noted in Section 3, shared trends are apparent across U.S. metropolitan areas, which may be as geographically heterogeneous as the nations in our sample. Presumably, if international markets are integrated in the long run, then tradable items like food and goods should share trends. More strongly, if factors of production are free to move across national boundaries in the long run (i.e., if the factor-price-equalization theorem of international economics applies), then even prices of services and rent may share trends. Tests of shared price trends across nations are complicated by fluctuating exchange rates between currencies. Our strategy is simply to include exchange-rate series as regressors in the first stage of the ADF2 test.

As usual, our null hypothesis is that there are no shared trends. We tilted the results (not shown) in favor of rejecting the null hypothesis by not adjusting the critical rejection values of the ADF2 test for the increase in the system dimension arising from the inclusion of the nuisance exchange-rate series. The main findings are briefly described. (The detailed results are available from us.) There were few cases in which the null hypothesis of no shared trends was rejected. There was no rejection for rent or goods. There was one for services and one for food (in each case between the United States and Canada). Although a shared price trend for the U.S.–Canada pair seems plausible, no shared trends were found between other plausible pairs such as Germany and the Netherlands. Overall, the paucity of shared trends is consistent with persistent barriers or wedges in trade between nations, which, for example, may arise from different trends in effective tax rates across nations. (It is interesting to speculate if the European Economic Community common-market consolidation of 1992 will lead to a new regime of more shared price trends, at least for European countries.)

5. CONCLUDING REMARKS

Statisticians and economists have long been fascinated with the behavior of prices. This article explores a hitherto neglected aspect of the subject—the behavior of (long-run) price trends and their interrelations. The degree of shared trends among prices is of considerable practical importance. For example, when agents have very different preferences, the usefulness of a universal price index can be justified by the presence of shared price trends. Similarly, the construction of national indexes is readily vindicated if prices of components share trends across regions. A finding of shared trends would also be consistent with the simplified models, such as the quantity theory of money, used by economists to explain or predict prices. In general, the number of explanatory factors in the determination of long-run prices cannot be less than the number of underlying unique trends.

Though our article focuses on price behavior and its implication for COLI's, the cointegration methods that we use may fruitfully apply to indexes used in other areas such as real estate (e.g., the price deflator for one-family structures prepared by the Bureau of Economic Analysis), school performance (school-district average Scholastic Aptitude Test scores), art (Rush's or Sotheby's), intelligence (the much-debated intelligence quotient), equity markets (the ubiquitous Dow–Jones Industrial Average), and economic activity (the Index of Industrial Production or the gross national product).

Our empirical findings are summarized in Table 9. Except for rents, there are shared trends among prices within metropolitan areas. Moreover, across areas within the United States, there appear to be arbitraging mechanisms for the prices of food, goods, and services.

Within OECD nations, we do not find shared price trends for any commodities, which suggests that more factors affect prices within these nations than within metropolitan areas of the United States. Looking at single-commodity groups across these nations, we find hardly any examples of shared trends, even between such close nations as Germany and the Netherlands; substantial impediments to the free flow of goods (or factors) may persist among these nations.

Employing the newly available techniques of cointegration, we have demonstrated some difficulties that defeat simple justifications of universal reliable indexes. Our results provide a starting point to explore a key pragmatic question: What level of aggregation properly trades off the loss of informativeness on distinct trends with the gain from reduction in data handling? A question that arises for understanding economies is: What

Table 9. Summary: Number of Distinct Trends in Prices

Components	*U.S. metropolitan areas*	*OECD nations*
	For four components within jurisdiction	
	3[a]	4[b]
	Across six jurisdictions	
Rent	6	6
Services	2	5[c]
Goods	2	6
Food	1	5[c]

NOTE: Only the number of trends that cannot be rejected as null hypothesis are included. Any greater number of distinct trends is rejected.

[a] For each of the six U.S. metropolitan areas, rent is a separate trend; food, goods, and services can be represented by two trends.

[b] Canada is the one exception, with food, goods, and services having only two distinct trends.

[c] Canada and the United States share a trend.

are the fundamental sources of the observed multiplicity of distinct price trends?

ACKNOWLEDGMENTS

The Japanese Corporate Associates Program of the John F. Kennedy School of Government provided research support. We thank James Stock and two anonymous referees for helpful comments.

[*Received March 1988. Revised May 1989.*]

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