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HICKSIAN AGGREGATION AND PRICE DYNAMICS: TESTS FOR A SINGLE WHEAT PRICE INDEX

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I. Introduction

Aggregation over commodities is a useful assumption in economic analysis. While some degree of aggregation is often inevitable because of the nature of data, this assumption is usually considered for empirical simplicity. In demand estimations, commodity aggregation greatly reduces the number of parameters to estimate. Spatial equilibrium or econometric simulation models also can be simplified with the assumption.

Wheat is a typical commodity for which aggregation is assumed over different varieties. Many data sources released by United States Department of Agriculture (USDA) report aggregate wheat price indexes (e.g., Agricultural Prices; Agricultural Statistics; Wheat Situation and Outlook Report) and voluminous studies deal with average price and quantity of wheat. However, data aggregation over different classes of wheat is questionable.

There are five different classes of wheat grown in the United States, the largest wheat exporter in the world: hard red winter (HRW), soft red winter (SRW), hard red spring (HRS), durum, and white wheat (WW). Each class of wheat is produced in a specific region with different growing conditions and has unique milling and baking properties (CBOT 1989). Some classes are processed for different end uses (Pomeranz 1988, 15) and, thus, do not belong to the

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same market in consumption[1]. In production, only a few classes of wheat compete for the same acreage (Briggle et al. 1982)[2]. These observations suggest that individual classes of wheat should be treated as separate commodities and that "average" wheat prices may misrepresent the price dynamics of individual wheat classes.

Varian (1992) discusses two cases under which aggregation over commodities is possible: 1) functional separability that imposes restrictions on the utility function, and 2) Hicks' aggregation that imposes restrictions on price dynamics (1956). Gorman (1959) show that a perfect price index exists if, and only if, the utility function is weakly separable into homogeneous subutility functions. The homothetic separability implies that commodities in the group is separable from all other commodities (or groups of) in the consumption bundle and that each commodity has the same income elasticity. This condition is intuitively unappealing and too strong for empirical data.[3] Further, few would assume this condition on the utility function and construct (should estimate for empirical applications) price indices from parameters of detailed data as the way the restriction implies, e.g., Divisia indices, to analyze data at a higher level of aggregation (Anderson 1979). Instead, most aggregate analyses use readily available average prices which impose restrictions on price dynamics of commodities in the group.

Hicksian aggregation suggests that if prices of goods in a group increase or decrease in the same proportion, a single price index can represent the prices; and the optimal amounts of the goods in the

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[1] HRW and HRS are used for bread, SRW and WW for cookies, cake, and snack foods, and durum for spaghetti and pasta products. The mixtures of HRW and HRS and of SRW and WW largely depend on the price of each class and product characteristics of end use.

[2] Winter wheat (HRW, SRW, and some WW) is planted in the fall and harvested in the summer, while spring wheat (HRS, durum, and some WW) is planted in the spring and harvested in the fall. HRW is mainly produced in Kansas, Nebraska, Oklahoma, and Texas and SRW in the midwest and on the Atlantic coast. HRS is grown mainly in northern states such as North and South Dakota, Montana, Minnesota, and Nebraska; and durum is grown primarily in North Dakota and Montana. WW is produced mainly in Michigan, New York, and the Pacific states.

[3] A representative model derived from a homothetically separable utility function is the Armington model, which has never been accepted by empirical data (e.g., Alston et al.).

[4] See also Varian (1983) for the proof of Hicks' aggregation, using the Generalized Axiom of Revealed Preference.

group can be solved in an aggregated form (Hicks' aggregation).[4] Proportional movements of prices imply "long-run co-movements" among the prices. Price ratios may deviate from the equilibrium (i.e., representative price) in the short run, but they should revert to a single price in the long run to guarantee the proportional changes.

Prices move together if they are close substitutes in demand and /or production. For instance, a stable long-run co-movement is likely between beef and pork prices or between different oil seed prices. This is because these goods are considered to be in the same market and, thus, to be governed by common factors that influence demand and supply, such as taste, recipes, nutrition ingredients, and technology.

The objective of this paper is to determine whether a single representative price index exists for wheat in the U.S. The U.S.data are examined since the country is the world largest wheat exporter and thus the U.S. price is most frequently used in many empirical studies. Testing existence of single price index involves a test of the necessary condition for the Hicksian aggregation, proportional price changes. The test used in this study is parametric, but, unlikely to parametric tests of functional separability, is not subject to parametric restrictions on the underlying utility or subutility functions (Blackorby, Primont and Russell 1970).

If the single price index in the wheat markets is not supported, the average price of the five classes of wheat would not correctly represent price dynamics of individual classes of wheat. Estimated demand elasticities or results from policy simulation models, using aggregate data, cannot represent market behaviors of individual wheat classes.

This paper is organized as follows. The next section explains economic conditions for Hicksian aggregation and develops test methods. The third section describes procedures and data. Then, results and interpretations are presented, followed by the section concluding this paper.

II. Hicksian Aggregation and Price Dynamics

Let $x=\{x_1, x_2, ..., x_n\}'$ be a vector of consumption goods and $p=\{p_1, p_2, ..., p_n\}$ be the corresponding price vector evolved from constant base prices, $p^0=\{p^0_1, p^0_2, ..., p^0_n\}$. Varian (1992) shows that if the vector p is proportional to the base price so that $p=kp^0$ for some scalar k, then a price index, $P=f(p)$, can be constructed to solve the optimal level of x in the aggregate form, $X=g(x)$. Specifically, the optimal solution for the standard utility maximization

[1] Max $U = U(x,z)$ subject to $y = px + qz$,

where y is the budget and q is the price of the other good z in the commodity bundle, is the same as that for

[2] Max $U = U(X,z)$ subject to $y = PX + qz$,

where $P=k$ and $X=p^0x$. If all prices in the price vector p move proportionally to the base prices in the same direction, then the price index P can be used to solve for the aggregate quantity index X, which behaves like ordinary quantity x.[5] The utility maximization problem is reduced to a manageable size with the assumption of Hicksian aggregation.

Under the assumption of proportional price changes, a price series can be expressed as

[3] $p_{it} = k_{it}p_i^0$ and $k_{it} = k_t$, for all $i = 1, ..., n$.

An arbitrary price ratio with the restriction is

[4] $p_{jt}/p_{1t} = k_tp_j^0/k_tp_1^0$, for all $j = 2, ..., n$,

or, equivalently,

[5] $\log(p_{jt})-\log(p_{1t}) = \log(c_j)$, for all $j = 2,..., n$,

where c_j is a constant. These imply that the (n-1) price ratios or differences in log prices should revert to constant means over time. The log difference can deviate from the steady state equilibrium, $\log(c_j)$, in the short run, but it must eventually revert to the equilibrium in the long run to ensure proportional price movements.[6]

The log difference would be stationary if each log price is stationary without differencing. However, stationary log prices alone do not guarantee proportional price changes because, from equation [3], it simply says that $\log(k_{it})$ is stationary. The condition of $k_{it}=k_t$ for all i is not yet confirmed.

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[5] Under Hicksian aggregation, P is equivalent to the simple average price deflated by the base-year average price.

[6] A key reason for empirical deviations from the equilibrium in the short run is "overshooting". Noisy or false information, inertia in consumer habit, and existence of contracts hinder prices from adjustusting instantaneously and cause temporary deviations, if they do not change fundamentals of the economy.

When each log price is stationary, the condition can be tested by estimating

[6] $\log(p_{jt}) = \alpha_j + \beta_j \log(p_{1t}) + \varepsilon_{jt}$, $j = 2, ..., n$,

where α_j and β_j are parameters and ε_{jt} is assumed i.i.d. normal with zero mean and finite variance. The null hypothesis of proportional price changes is Ho: $\beta_j=1$ for all j.[7]

When prices under consideration are nonstationary, however, the test procedure above would not be valid.[8] Granger and Newbold (1974) point out that the ordinary regression of two nonstationary variables may lead to spurious estimates and test results. However, nonstationary prices do not immediately rule out a stable long-run equilibrium between prices. Even though individual time series are nonstationary, a linear combination of series can be stationary. This property of time series data is called "cointegration" (Granger 1981).

The log difference in equation [5] can be stationary if the log prices are cointegrated with the cointegration vector (1 -1). The stationary log difference implies that $\log(k_{jt})-\log(k_{1t})$ converges to zero and $k_{jt}=k_{1t}$ in the long run.

The cointegration test for proportional price movements is a joint hypothesis test, i.e., the cointegration relationship should hold for all (n-1) pairs in the price vector. However, statistics for the joint cointegration test are not available in the literature. An alternative test for the joint hypothesis is to test the long-run relationship among n prices. If all (n-1) pairs in the vector are cointegrated, the n prices would also have, at least, a stable equilibrium.

To determine the relationship among the prices in the group, equation [5] is summed over j as

[7] $\log(p_{2t}) + ... + \log(p_{nt}) - (n-1)\log(p_{1t}) = \Sigma_j \log(c_j)$.

When individual prices are not stationary, the assumption of Hicksian aggregation requires a stable long-run relationship among log prices in the group. It should be noted, however, that equation [7] does not specify the cointegration vector because the vector is subject to normalization. Without the long-run equilibrium, each price has its own dynamic path; and the prerequisite for Hicksian aggregation does not hold.

[7] The OLS estimate of β_j reduces to unity under the assumption of proportional price changes in equation [3].

[8] Past research suggests that nonstationarity is difficult to reject for many economic time series, including commodity prices (e.g., Ardeni 1989; Goodwin 1992).

III. Test Procedures and Data

In the previous section, we have shown that the validity of the Hicksian aggregation can be checked with simple regressions if prices are stationary or with cointegration tests if prices are nonstationary. Thus, a unit root test is a prerequisite for the test for proportional price changes. This section discusses procedures for the unit root and cointegration tests and describes the data for empirical analysis.

UNIT ROOT TEST

In the literature, numerous tests of a unit root have been suggested. Engle and Granger (1987) summarize seven test procedures of nonstationarity. While most tests are based on i.i.d. normal, zero mean errors, Phillips and Perron (1988) developed a test statistic that allows weakly dependent and heterogeneously distributed innovations. Yang and Brorsen (1992) found evidence of dependence and conditional heteroskedasticity in some daily wheat price changes. This study uses the Phillips-Perron test. For the test, regress

[8] $\ln(p_{it}) = \alpha 1n(p_{it-1}) + e_t$, for all $i = 1, \ldots, n$,

and compute

[9] $Z_t = (2\Sigma_t \ln(p_{it-1})^{0.5}(\alpha-1)/s_T - .5(s_T^2-s_o^2)[T^{-1}s_T^2(2\Sigma_t \ln(p_{it-1}))^{0.5}]^{-1}$,

where T is the sample size, $s_o^2=T^{-1}\Sigma_t e_t^2$, is the consistent variance estimate under the null of a unit root, and $s_T^2=T^{-1}\Sigma_t e_t^2+2T^{-1}\Sigma_q\Sigma_{t=q+1}e_t e_{t-q}$, is the consistent variance estimate in the presence of weakly dependent and heterogeneously distributed errors. If this statistic is greater than the critical values in Fuller (1976, p. 373), the null hypothesis of presence of a unit root in $\ln(p_i)$ is rejected.

COINTEGRATION TEST

Tests for cointegration have been developed by Engle and Granger (1987) and Johansen (1991). The Granger-Engle method is simple. However, empirical results from this method often are sensitive to normalization (Johnson and Yang 1991). The Johansen method has become the convention for multivariate tests and is used in this study.

Let Y_t be an (nx1) vector of variables of interest, i.e., $\ln(p_t)$ in this study, which can be specified as a VAR form:

[10] $\triangle Y_t = \Gamma_1 \triangle Y_{t-1} +\ldots+ \Gamma_{q-1}\triangle Y_{t-q+1}+\pi Y_{t-1}+\mu_0+\delta D_t+\varepsilon_t$,

where μ_0 is the intercept, D is a vector of monthly dummies capturing deterministic seasonality, and ε is an n-dimensional vector of i.i.d. normal errors with mean zero. The rank of π (say, r) indicates the number of cointegration relationships among the variables in Y_t. If r=n, each time series is stationary in level. While r=0 indicates no long-run equilibrium among the variables, 0<r<n indicates r stable relationships among the variables. Thus, the test for a cointegration relationship is equivalent to a test for Ho:r=0.

Two auxiliary regressions are needed for the test:

[11] $\Delta Y_t = \Gamma_{11}\Delta Y_{t-1} + \ldots + \Gamma_{1q-1}\Delta Y_{t-q+1} + \mu_{10} + \Psi_1 D_t + \varepsilon_{1t}$,

[12] $Y_{t-q} = \Gamma_{21}\Delta Y_{t-1} + \ldots + \Gamma_{2q-1}\Delta Y_{t-q+1} + \mu_{20} + \Psi_2 D_t + \varepsilon_{2t}$.

The residual vectors, ε_{1t} and ε_{2t}, are used to calculate squared canonical correlations, $\lambda_1 > \lambda_2 > \ldots > \lambda_n$. The test statistics are computed, using these correlation coefficients.

Johansen (1991) developed two test statistics. The trace test determines the null hypothesis that there are, at most, rcointegration relationships, while the maximum eigenvalue test determines the null hypothesis of r cointegration vectors against r+1 cointegration vectors. The maximum eigenvalue test would be more powerful than the trace test since the former tests against a specific alternative hypothesis (Johansen and Juselius 1990).

The maximum eigenvalue test statistic for r cointegration vectors is

[13] $\mathrm{MET} = -T\ln(1-\lambda_{r+1})$, r = 0, 1, ..., n-1.

Critical values for equation [10] are tabulated in Osterwald-Lenum (1991). He generated critical values under the null hypothesis of r cointegration vectors independently of those under the null hypothesis of r' cointegration vectors, where $r' \neq r$.[9]

DATA DESCRIPTION

Data used for this study are monthly prices of five classes of wheat in the U.S. markets: HRW, SRW, HRS, WW, and durum. The sample period is January 1981 through December 1990 (10 years). For more general conclusions, three different sets of the five prices

[9] The Johansen method may not be robust to conditional heteroskedasticity, which is often found in the distributions of commodity prices. A solution for this is to generate critical values adjusted for conditional heteroskedasticity. This statistical work is beyond the current study.

are composed over marketing stages: farm, domestic, and export prices. The farm and domestic prices are taken from the 1993 USDA Wheat Situation and Outlook Report, while the export prices are from various issues of International Wheat Council's World Wheat Statistics.

Specifications for each price are reported in Table 1. For farm prices, the data sources do not report durum prices separately. The Northern Plains spring wheat price is an aggregated price of durum and HRS wheat. For some months, durum export prices have not been quoted due to negligible or no trade. For the data, the previously quoted prices are used under the assumption that the previous prices still represent the market.

TABLE 1 Specifications of price data, Jan. 1981 - Dec. 1990

	Farm Prices
HRW	Central and Southern Plains
SRW	Corn Belt
HRS	Northern Plains
WW	Pacific Northwest
	Domestic Prices
HRW	Kansas City, No. 1 (13% protein)
SRW	Chicago, No. 2
HRS	Minneapolis, No. 1 (13% protein)
WW	Portland, No. 1
Durum	Minneapolis, No. 1
	Export Prices (fob)
HRW	Gulf, No. 2 (ordinary protein)
SRW	Gulf, No. 2
HRS	Pacific, No. 2 (14% protein)
WW	Pacific, No. 2
Durum	Lakes, No. 3

Note: Durum farm prices are not reported separately in the data source. The HRS price received by farmers is an aggregate price of hard red spring and durum wheats. Units are \$/60-pound bushel for farm prices and \$/bushel for domestic and export prices.

IV. Empirical Results and Interpretations

Table 2 shows test results for unit roots in log prices. For all prices, the null hypothesis of a unit root in level is not rejected at conventional levels of significance. This implies that log prices at different marketing levels are not mean reverting, but wander without bounds. The hypothesis of proportional price changes is, therefore, tested with the cointegration test.

TABLE 2 Phillips-Perron unit root test results

Prices	Levels	First differences
	Farm prices	
HRW	-0.22	-62.92*
SRW	-0.26	-58.66*
HRS	-0.16	-51.89*
WW	-0.22	-63.18*
	Domestic prices	
HRW	-0.16	-90.32*
SRW	-0.36	-91.93*
HRS	-0.26	-77.86*
WW	-0.19	-77.88*
Durum	-0.40	-85.28*
	Export prices	
HRW	-1.78	-141.09*
SRW	-2.31	-114.53*
HRS	-1.77	-119.00*
WW	-2.34	-107.48*
Durum	-2.02	-106.38*

Note: Critical values are tabulated in Fuller. For the sample size 100, they are -2.89 at the 5% level and -3.50 at the 1% level. One asterisk indicates rejection of the null hypothesis of a unit root at the 1% significance level

Table 3 shows the cointegration test results at the three different marketing levels. All three MET statistics for r=0 are less than the critical values at a 5% level. This implies that the null hypothesis of no cointegration relationship is not rejected against the alternative hypothesis of one cointegration relationship. A vector of time series may be cointegrated of order d>1. However, this is not the case for the prices in this study since the first differences of the prices are all stationary at a 1% significance level (Table 2). The hypothesis of a long-run equilibrium among the five different wheat classes is not supported at any marketing level.

TABLE 3 Maximum eigenvalue test results for five wheat prices, Jan. 1981 - Dec. 1990

Null hypothesis[a]	MET	Critical 5%	value[b] 10%
	Farm prices (lag=1)[c]		
No cointegration	20.25	27.07	24.73
1 cointegration	11.61	20.97	18.60
2 cointegrations	8.57	14.07	12.07
3 cointegrations	0.44	3.76	2.69
	Domestic prices (lag=1)		
No cointegration	28.59	33.46	30.90
1 cointegration	17.44	27.07	24.73
2 cointegrations	12.97	20.97	18.60
3 cointegrations	9.50	14.07	12.07
4 cointegrations	3.12	3.76	2.69
	Export prices (lag=1)		
No cointegration	22.56	33.46	30.90
1 cointegration	16.32	27.07	24.73
2 cointegrations	11.98	20.97	18.60
3 cointegrations	6.06	14.07	12.07
4 cointegrations	1.08	3.76	2.69

[a] If the null hypothesis is r cointegration vectors, the alternative hypothesis of the maximum eigenvalue test is r+1 cointegration vectors

[b] Critical values are taken from Osterwald-Lenum.

[c] The lags are in levels. Akaike's information criterion is used.

The cointegration test results provide evidence against proportional movements of wheat prices in the U.S. markets. Thus, current wheat price indices would misrepresent individual wheat price dynamics. Given consumption and production circumstances, these results are not surprising since the five different classes of wheat have no single market.

A question then would be "Is there any long-run equilibrium among individual wheat prices whose products compete (or substitute) in production or end use?" This is an empirical question and can be answered by pretesting. However, with five prices, 26 (${}_5C_5 + {}_5C_4 + {}_5C_3 + {}_5C_2$) different subgroups are involved, including the case examined above. Thus, the five prices are divided into subgroups, based on a priori information about competition in production and end use.

In end use, HRW and HRS are processed mainly for bread; and durum is used exclusively for pasta products, while a small amount of HRS flour is added to semolina to make spaghetti or pasta products. SRW and WW are used mainly for cookies and cake. Hard wheat is added to soft wheat for noodles, especially in Asian countries, but the ratio of hard wheat to the total mixture is negligible. In production, HRS and durum compete for limited land in durum-producing regions, and HRS and HRW compete in HRS producing regions other than North Dakota. Thus, the five wheat classes are divided into two subgroups: one for HRW, HRS, and durum and the other for SRW and WW.

Cointegration test results for these two subgroups are shown in Tables 4 and 5. The MET statistics for no cointegration among HRW, HRS, and durum prices is rejected at all three marketing levels (Table 4). At domestic levels, two stable equilibria exist. These results support our proposition that a long-run equilibrium exists for prices in the same market.

On the other hand, the null hypothesis of no cointegration between SRW and WW prices is not rejected (Table 5). The MET statistics for the three marketing levels are all less than the critical value at the 5% level. However, this does not contradict our proposition. Even though SRW and WW have the same end use, their producing regions are geographically separated with SRW in midwest and Atlantic Coast states and WW in Pacific Coast states. Because of

TABLE 4 Maximum eigenvalue test results for HRW, HRS, and durum prices, Jan. 1981 - Dec. 1990

Null hypothesis[a]	MET	Critical 5%	value[b] 10%
	Farm prices (lag=2)[c]		
No cointegration	14.85	14.07	12.07
1 cointegration	2.49	3.76	2.69
	Domestic prices (lag=1)		
No cointegration	22.02	20.97	18.60
1 cointegration	15.85	14.07	12.07
2 cointegrations	3.37	3.76	2.69
	Export prices (lag=2)		
No cointegration	32.60	20.97	18.60
1 cointegration	9.90	14.07	12.07
2 cointegrations	2.14	3.76	2.69

[a] If the null hypothesis is r cointegration vectors, the alternative hypothesis of the maximum eigenvalue test is r+1 cointegration vectors.

[b] Critical values are taken from Osterwald-Lenum.

[c] The lags are in levels. Akaike's information criterion is used.

distance and transportation costs, they hardly compete for the same geographic markets in the U.S. In exports, WW is shipped mostly to Asia, while SRW is shipped mostly to Europe and Africa. We do not expect any strong co-movement among prices in separate markets.[10]

[10] This can explain poor performance of hedging white wheat at the Chicago Board of Trade, where SRW is the dominating class traded.

TABLE 5 Maximum eigenvalue test results for SRW and WW prices, Jan. 1981 - Dec. 1990

Null hypothesis[a]	MET	Critical 5%	value[b] 10%
	Farm prices (lag=2)[c]		
No cointegration	0.56	14.07	12.07
1 cointegration	3.62	3.76	2.69
	Domestic prices (lag=2)		
No cointegration	10.52	14.07	12.07
1 cointegration	3.04	3.76	2.69
	Export prices (lag=2)		
No cointegration	11.27	14.07	12.07
1 cointegration	2.68	3.76	2.69

[a] If the null hypothesis is r cointegration vectors, the alternative hypothesis of the maximum eigenvalue test is r+1 cointegration vectors.

[b] Critical values are taken from Osterwald-Lenum.

[c] The lags are in levels. Akaike's information criterion is used.

V. Conclusions

Wheat is not simply wheat. For instance, durum wheat, which is produced in the Northern Plains and processed into pasta products, should be treated separately from white wheat, which is produced in Pacific states and used for cookies. The empirical tests in this study provide evidence against the assumption of Hicksian aggregation across five wheat classes in the U.S.markets. The results imply that there is no such a price index legitimately representing five classes of wheat in the U.S. markets.

Empirical analysis using statistical average wheat prices would suffer from aggregation bias. The magnitude of bias should depend on characteristics of research, such as data, model specifications and/or objectives. However, consequences of aggregation over five wheat classes go beyond what econometric criteria can tell.

Price elasticities of wheat, based on the aggregate data, may not provide reliable information to producers or research of individual classes of wheat. Policy simulation models dealing with aggregate wheat data may not necessarily be applicable to individual classes of wheat markets. The heterogeneous price dynamics found in this study also suggest that agricultural price policies commonly applied to all classes of wheat (e.g., target price, loan rate, and EEP bonus) would have differential impacts on those who operate in these separate markets. Depending upon the goals of and political constraints facing policy-makers in establishing these policies, considerations of differentiating the classes would be appropriate.

This study is in the same line as Leontief's (1993) concerns about data aggregation:

> Needless to say, results of empirical computation can be expected to yield different results, depending on the use of any of the possible alternative classifications. The only reliable method of assessing the validity of these computations would be to avoid combining essentially heterogeneous elements, that is, to use a more detailed classification.

The results in this study suggest using disaggregate wheat data for more sensible results. Data disaggregation may introduce complexity to research. This, however, can be alleviated with the rapid progress in computer software as well as hardware.

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