HAVE PRICE LINKAGES IN THE WORLD MARKET OF COTTON IMPROVED OVER THE LAST DECADE?

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Abstract

This paper examines the degree to which cotton prices are linked and also tests whether such price linkages have improved over the last decade. It concludes that while some cotton markets are well-linked others are not. The degree of linkage has improved over the last decade while the main source of this improvement appears to be a result of short-run price transmission rather than long-run comovement.

Introduction

One of the most vital elements of international trade theory is that the price of a particular commodity in different locations is the same. However, because of transportation costs, differences in quality, trade restrictions, and different times at which the agreement to trade and the actual transaction takes place, prices are unlikely to be the same. However, one does expect some degree of linkage. In other words, prices changes in one region are expected to be followed by similar price changes in other regions where the same commodity is traded. The existence of strong price linkages ensures that, among others, resources are allocated in an efficient manner.

The objective of this paper is to examine: (a) how strong are the price linkages within the world cotton market and (b) whether such linkages have improved over the last decade. In pursuing these two objectives, the present paper contributes to literature of price linkages in two respects. On the theoretical side, it introduces a well-defined measure of price linkage by allowing one to identify the source of such linkages (i.e. whether improvement is a result of short-run price transmission rather or long-run comovement.) On the empirical side it applies this measure to the world market of cotton for two different time periods thereby examining whether improvement in price linkages has taken place.

There are two reasons as to why one would expect that price linkages may have improved over the last decade. First, advances in information technology have made it much easier for information on domestic demand/supply conditions of cotton (or any other commodity) to be disseminated across countries. Second, recently a number of countries have undertaken steps to liberalize their cotton subsectors and thus eliminating impediments to trade.

The remainder of the paper proceeds as follows. In the next section we develop the model which measures the degree of market linkages in the world market of cotton. In the third section we describe the data and discuss the results while the last section concludes.

Detecting Price Linkages

Earlier studies examining the relationship between set of prices either have looked at correlation coefficients [e.g., Timmer, Falcon, and Pearson (1983); Stigler and Sherwin (1985)] or have used the following regression [e.g. Isard (1977), Richardson (1978), Mundlak and Larson (1992)]:

$$p_t^1 = \mu + \beta_1 p_t^2 + \varepsilon_t , \tag{1}$$

where p_t^1 and p_t^2 denote prices from two origins of the commodity under consideration, μ and β_1 are parameters to be estimated while ϵ_t denotes an i.i.d ~ N(0, σ^2) error term. The hypothesis that the slope coefficient equals unity and (possibly) the intercept term equals zero can be tested; formally, H_0: $\mu + 1 = \beta_1 = 1$. Under H_0 the deterministic part of (1) becomes $p_t^1 = p_t^2$, in turn implying that the price differential, $p_t^1 - p_t^2$, is an i.i.d ~ N(0, σ^2) term.

Estimating (1) and testing H_0, while intuitively appealing and computationally implementable, presents two fundamental shortcomings. First, in primary commodity markets factors such as (small or even perceived) differences in quality, high transportation costs relative to the price, etc., it is rather unlikely that the two prices will only differ by an i.i.d ~ N(0, σ^2) term as H_0 of (1) dictates. Therefore, H_0 is expected to be rejected without necessarily ruling out a relatively high degree of linkage between the two prices.

Second, some statistical properties of the series involved in the regression, namely nonstationarity, may invalidate standard econometric tests and thus give misleading results regarding the degree to which price signals are being transmitted from one market to another. Consequently, it is deemed necessary to employ a model, general enough, that first relaxes the restrictive nature of (1) and second imposes no *a priori* requirements on the stationarity properties of the variables in question.

With respect to the nonstationarity problem one can examine the order of integration of the error term in (1) and make inferences regarding the validity of the model (Ardeni, 1989). If prices are indeed nonstationary, the existence of a stationary error term implies comovement between the two prices. However, if the slope coefficient is different from unity, the corresponding price differential would be growing

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and such growth would not be accounted for, although prices move in seemingly synchronous manner. Hence, stationarity of the error term of (1) (given non-stationary prices) while establishing proportional price movement, should not be considered as a testable form equivalent to that of the H_0 of (1).

To account for the non-unity slope coefficient one can restrict the parameters of (1) according to H_0. In such case, the problem is equivalent to testing for unit root in the following univariate process (Engle and Yoo, 1987):

$$(p_t^1 - p_t^2) \sim I(0). \tag{2}$$

If the price differential as defined in (2) is stationary, then one can conclude that price signals are transmitted from one market to another, in the long run. The assumption (or finding) that the cointegration parameter is unity is very crucial, as it ensures that there is no other nonstationary component entering the system. As Meese (1986) and West (1987) observe, the absence of cointegration (with unity slope coefficient in this case) can be attributed to omitted nonstationary variables, in turn implying that an additional component would have to be included in (2) in order to fully account for the variability of the price differential.

As a sidelight, it should be emphasized that if the cointegration parameter is unity, it is immaterial for all relevant aspects of the analysis whether (1) or (2) is employed. This is the case because as the sample size increases, regression (1) should yield β equal to unity. However, in finite samples this may not be necessarily the case. For example, Ardeni (1989), using (1) in logarithms for a number of internationally traded primary commodities, found that the corresponding error term was not stationary, thus rejecting the law of one price. Baffes (1991), on the other hand, by using the same data set found that in the majority of cases the price differential was stationary, hence providing supportive evidence for the law of one price.

From the preceding discussion, it is rather evident that cointegration tests are not very powerful as they only make inferences about the existence of the moments of the distribution of $(p_t^1 - p_t^2)$ and not about certain restrictions that may be required by economic theory [*e.g.*, H_0 of (1)]. Therefore, (2) cannot serve as a substitute for the H_0 of (1); it can only serve as an intermediate step in establishing its validity.

The restrictive nature of (1) can be circumvented by extending it to a more general autoregressive structure. Introducing one lag to (1), gives:

$$p_t^1 = \mu + \beta_1 p_t^2 + \beta_2 p_{t-1}^2 + \beta_3 p_{t-1}^1 + u_t, \tag{3}$$

Different restrictions on the parameter space of (3) result in different models. Hendry, Pagan, and Sargan (1983) discuss a number of testable hypothesis, results of corresponding restrictions on (3). An important one, applicable to a variety of economic models, is the long-run proportionality or homogeneity hypothesis: it ensures that price movements in one market (p_t^2) will *eventually* be transmitted to the prices of the other market (p_t^1). Such hypothesis can be tested by restricting all slope parameters of (3) to sum to unity (i.e., $\Sigma_i\beta_i = 1$).

Imposing long-run proportionality (i.e. setting $\beta_2 = 1 - \beta_1 - \beta_3$) in (3) and rearranging terms results in:

$$(p_t^1 - p_{t-1}^1) = \mu + (1-\beta_3)(p_{t-1}^2 - p_{t-1}^1) + \beta_1 (p_t^2 - p_{t-1}^2) + u_t. \tag{4}$$

Relationship (4) belongs to the family of error correction models (ECM). Because of the equivalence of the existence of cointegration and ECM specification (Engle and Granger, 1987), stationarity of the price differential implies (4) (specifically that $(1 - \beta_3)$ is significantly different from zero) and vice-versa.

The main feature of (4) (or alternatively (3)) is the economic interpretation of its parameters: β_1 indicates how much of a given change in the price of the commodity within the first period will be transmitted to the other price (referred to as initial adjustment, short-run effect, or contemporaneous effect); $(1 - \beta_3)$ indicates how much of the price difference between the two prices is eliminated in each period thereafter (referred to as error-correction, speed of adjustment, or feedback effect). The coefficient of the short-run effect can, in theory, take any value. The adjustment coefficient, however, is restricted between zero and one. The closer to unity is $(1 - \beta_3)$, the higher the speed at which convergence will take place. Long-run convergence requires $(1 - \beta_3)$ to be significantly different from zero; that is, $(1 - \beta_3)$ different from zero is a necessary and sufficient condition for long-run convergence. On the other hand, significantly different from zero β_1 is neither a sufficient nor a necessary condition for long-run proportionality; note that even if $\beta_1 = 1$ the series may drift apart in the long run – unless $(1 - \beta_3)$ is significantly different from zero, in which case the series will converge even if $\beta_1 = 0$.

The model outlined above suggests that, given that long-run proportionality exists, whether to choose (3) or (4) in order to recover short- and long-run dynamic price behavior is a matter of stationarity properties. If prices are stationary, (3) would be the preferred structure and long-run proportionality can be tested by restricting the parameters to sum to unity. Under non-stationarity, (4) would be the preferred structure and long-run proportionality can be tested by examining the stationarity properties of the price differential (Engle and Yoo, 1987) or equivalently testing whether $(1-\beta_3)$ is different from zero (Phillips and Loretan, 1991).

Having established long-run proportionality and also having recovered the parameter estimates of (4) (or the restricted form of (3)) the next task is to transform the information contained in the parameter space in such a way so that a succinct interpretation of both short-run and feedback effect (and hence price linkage) can be given. Stating the question otherwise: *How long does it take for the price of the commodity from origin 1 to adjust to a given price change in origin 2?*

Let n be the period in which k percent of the cumulative adjustment takes place. In the current period, $n = 0$, k takes the value of β_1 [also equal to $1\text{-}(1\text{-}\beta_1)$], which is the short-run impact of $(p_t^2 - p_{t-1}^2)$ on $(p_t^1 - p_{t-1}^1)$. In the next period, $n = 1$, k takes the value of $\beta_1+(1\text{-}\beta_1)\beta_3$, which is the impact of the previous period, β_1, plus the feedback effect, $(1\text{-}\beta_1)\beta_3$ [it can also be written as $1\text{-}(1\text{-}\beta_1)(1\text{-}\beta_3)$]. For $n = 2$, k takes the value of the previous period, $\beta_1+(1\text{-}\beta_1)\beta_3$ plus $(1 - \beta_3)(1\text{-}\beta_1\text{-}(1\text{-}\beta_1)\beta_3)$ [which can be written as $1\text{-}(1\text{-}\beta_1)(1\text{-}2\beta_3+\beta_3^2)$ or $1\text{-}(1\text{-}\beta_1)\beta_3^2$]. The following table gives the adjustment for the first four periods.

Period	**Amount of Cumulative Adjustment**	
0	β_1	$= 1 - (1 - \beta_1)\beta_3^0$
1	$\beta_1 + (1 - \beta_1)(1 - \beta_3)$	$= 1 - (1 - \beta_1)\beta_3^1$
2	$1 - (1 - \beta_1)\beta_3 + (1 - \beta_3)(1 - \beta_1)\beta_3$	$= 1 - (1 - \beta_1)\beta_3^2$
3	$1 - (1 - \beta_1)\beta_3^2 + (1 - \beta_3)(1 - \beta_1)\beta_3^2$	$= 1 - (1 - \beta_1)\beta_3^3$
4	$1 - (1 - \beta_1)\beta_3^3 + (1 - \beta_3)(1 - \beta_1)\beta_3^3$	$= 1 - (1 - \beta_1)\beta_3^3$

Hence, the cumulative adjustment at period n is given by:

$$k = 1 - (1 - \beta_1)\beta_3^n \tag{5}$$

For values of β_1 and β_3 close to unity, a small n (number of periods) is required for the adjustment to be completed (*i.e.* k close to unity). Alternatively, solving for n in (5) gives the cumulative adjustment achieved in n periods, i.e. $n = [log(1\text{-}k) - log(1\text{-}\beta_1]/log\beta_3$.

Data and Results

Data

Weekly (Thursday) quotations reported by Cotton Outlook are used in the analysis. All quotations (expressed in US ¢/lb.) are CIF, North Europe, cash against documents on arrival of vessel, including profit and agent's commission. Two samples, one covering the period August 15, 1985 to December 24, 1987 (122 observations) and a second covering the period August 3, 1995 to January 9, 1997 (73 observations) were constructed. A total of four cotton price quotations from the following origins were used: US (Memphis Territory), Greece, Central Asia, and African 'Franc Zone' (referred to as W. Africa). In addition to the four quotations, we also included the A Cotlook Index – a measure of the 'world' price of cotton.

The 'world' price of cotton (generally referred to as the Cotlook A Index or the A Index) is an index constructed daily by the Cotlook Limited, a private information dissemination company based in Liverpool, UK and is published in the weekly magazine *Cotton Outlook*. It is a simple average of the 5 less expensive styles of cotton (Middling 1-3/32'') out of are 14 cotton styles traded in North Europe. The four prices used in this study are used for the construction of the A Index; however, they may or may not be used depending on whether they are part of the 5 less expensive ones.

Since not all countries actively participate in the cotton export market throughout the year only the price series for which adequate sample length was available were examined.

Stationarity Tests

To determine the order of integration the augmented Dickey-Fuller (ADF) and the Phillips-Perron (PP) procedures were utilized. The ADF (Dickey and Fuller, 1981) is based on the following regression: $(x_t - x_{t-1}) = \mu + \beta x_{t-1} + lags(x_t - x_{t-1}) + \epsilon_t$ where x_t denotes the series under consideration. A negative and significantly different from zero value of β indicates that x_t is $I(0)$. The PP test (Phillips and Perron, 1988; Phillips, 1989) is similar to the ADF; their difference lies on the treatment of any nuisance serial correlation aside from that generated by the hypothesized unit root. To identify the presence of one unit root we test H_0: x_t *is not* $I(0)$ <u>against</u> H_1: x_t *is* $I(0)$. Note also that the significance level of the error-correction coefficient, β_3, can serve as cointegration test.

The upper panel of Table I reports stationarity results in levels for both periods. The tests indicate that stationarity in levels is rejected in all cases. The middle panel of Table I reports results for trend stationarity tests. Here the picture changes considerably since in all cases both ADF and PP statistics improve (i.e. become larger in absolute terms) and also in the second period, with the exception of the ADF statistic for C. Asia, all tests indicate that the prices are trend stationary and in some cases the evidence is very strong. Since it is unlikely that the order of integration of the price series has changed throughout the sample period, such result should be attributed to the low power of stationarity tests.

The lower panel of Table I reports stationarity statistics of the price differential. This gives a measure of the degree of comovement between pairs of cotton prices. Consider first the A index. When compared to the US in period 1 no comovement appears to be in place, while a high degree of comovement is present in the second period (both tests are consistent). A similar result holds for A Index-Greece, where the level of significance increases from 5 and 10% in the first period to 1 and 5% in the second period. The link between the A Index and the remaining two prices, however, appears to be weakening in period 2. The A Index-W. Africa price differential, while stationary in period 1, it is non-stationary in period 2.

The degree of comovement of prices increased substantially in Greece, W. Africa, and C. Asia, when coupled with the US. In most cases, stationarity statistics more than doubled and in all but one case they exceeded the 5% significance level. Comparing Greece with W. Africa and C. Asia, the exact reverse is true. In both cases, the comovement sharply deteriorates. Finally, for W. Africa - C. Asia, while the statistics become lower in absolute value, they are still significant at the 5 and 10% level.

To conclude, results from the lower panel of Table I indicate that, excluding the A Index, price linkages in the cotton market improved relative to the US but no improvement was detected among non-US markets. These results are robust with respect to both stationarity tests (PP and ADF).

Goodness of Fit

The two samples were estimated together and a Chow test was used to determine whether the parameters of period 1 were significantly different from those of period 2 while the χ^2 testing procedure proposed by Hansen (1982) and White (1980) was utilized to estimate the covariance matrix consistently.

To assess the overall performance of the model, we first examine the goodness of fit (reported in Table II). Given that (4) (and also the restricted version of (3)) can be reparameterized in terms of current and lagged price differentials as well as one of the two price differences – also current and lagged – (Campbell and Shiller, 1987), one can think of the R^2 as a measure of basis risk (the unpredictable movements in the basis) where basis is defined as the difference between the two prices (rather than its usual definition as the difference between cash and futures). Then, the lower the R^2 the higher the basis risk and vice-versa.

With the exception of a marginal reduction in the A Index-C. Asia case (from 0.88 to 0.87), the R^2 has improved substantially in all remaining cases. On average, about 50% of the variability of price in one region was explained by the variability of another region's price in period 1. In period 2 the explanatory power of the model increased to 75%. Thus, with the evidence at hand, it appears that price linkages within cotton markets have improved substantially over the last decade. The next step is to examine whether such a conclusion can be deduced if further measures of are applied and also identify the sources of such improvement.

Quantifying the Improvement in Price Linkages

Table III reports the short-run effect or β_1 of specification (4). The upper and middle panels depict the adjustment, which takes place within the first period, due to an exogenous change in the prices in another region in sub-sample periods one and two respectively. Thus a coefficient of one would be interpreted as a perfect transmission of price shocks in another region, while a coefficient of zero represents an invariance of the prices to changes in prices elsewhere. Since the short-run effect is not restricted to numbers between 0 and 1, a β_1>1, for example, would suggest an over reaction to changes in prices in the current period. The lower panel contains the *p*-values resulting from the test of the hypothesis of equality in the β_1s in the two sub-sample periods, against the two-sided alternative.

At the 5% significance level, six of the nine overall improvements in the short-run effect were significant, while only three of the eight cases represented significant reductions in the amount of adjustment within the first period. Further analysis of the nine significant changes in the short-run effect between the two periods revels that the average deviation of the adjustment coefficient from one fell from 0.32 to 0.25. The above result indicates an overall improvement in the immediate speed of adjustment, in response to a given change in prices in another region.

Greece, by far, showed the most improvement in the short-run adjustment when coupled to the A Index, US and C. Asia. However, W. Africa and C. Asia revealed signs of improvement when paired with Greece, while the opposite was true when paired with the US.

The measure of long-run comovement in the cotton prices (i.e., (1 - β_3) in specification (4)) is presented in Table IV, with the upper and middle panels representing the effect in period 1 and 2 respectively. In essence, the measure of long-run adjustment captures the correction to a given price change in another region subsequent to the current period. In fact, the absolute deviation from the long-run steady-state declines from period to period (i.e., suggesting long-run comovement in prices) when this parameter is statistically significant. The lower panel of Table IV reports the *p*-values for the test of the hypothesis that the dynamic adjustment effect remained the same against the two-sided alternative.

Twelve improvements were observed, while declines in the degree of comovement were present in three cases between the two sub-sample periods. The remaining five cases revealed no appreciable change between periods 1 and 2. Significant improvements in the long-run effect were observed when Greece was coupled with A Index and W. Africa at the 6% significance level. However, all other changes in the measure of long-run comovement between the two sub-sample periods appear to be insignificant at conventional levels of significance.

Table V presents the number of weeks, *n*, required to achieve 95% of the adjustment to a given price change. Note that *n* is calculated using equation (5) and it is only meaningful when long-run comovement, in the Engle-Granger sense, is detected. Faster adjustment is observed in fourteen cases while a slower adjustment period is observed in only two. Except Greece-US in period 2, with the US as a reference, it is clear that none of the other regions exhibited convergence towards the price levels in the US. This fact becomes apparent when the insignificant error-correction coefficient reported in Table V is considered.

However, Table VI reveals that nine of fourteen changes in the number of periods required to achieve 95% of the adjustment, were significant at the 7% significance level. Hence, price shocks were transmitted at higher speed in period 2 compared with period 1. In addition, in period 1, nine cases of non-convergence, were evident, while only three cases appeared in period 2 at the 10% level of significance. The above observations indicate that more regions achieved long-run comovement in the second period.

TABLE I: Stationarity Tests for Price Levels (w/o and w/ trend) and Price Differentials

	Period 1		***Period 2***	
	ADF	*PP*	*ADF*	*PP*
Levels w/o trend				
A Index	-1.24	-0.70	-1.18	-0.84
US	-1.40	-1.10	-1.38	-1.53
Greece	-1.26	-0.78	-1.18	-0.96
W. Africa	-1.33	-0.72	-1.11	-0.67
C. Asia	-1.30	-0.75	-1.24	-0.86
Levels w/ trend				
A Index	-2.42	-2.03	-	-
US	-1.83	-1.55	-	-
Greece	-2.55	-2.27	-2.87*	-
W. Africa	-2.56	-2.07	-	-
C. Asia	-2.80*	-2.35	-2.38	-2.59*
Price Differentials				
A Index - US	-1.71	-1.40	-	-
A Index -	-	-2.82*	-	-
A Index - W.	-	-	-2.48	-2.41
A Index - C.	-	-	-	-2.65*
US - Greece	-1.87	-1.71	-2.81*	-
US - W. Africa	-1.74	-1.46	-	-
US - C. Asia	-1.68	-1.28	-	-
Greece - W.	-	-2.96**	-2.49	-2.35
Greece - C.	-2.85*	-2.80*	-2.26	-2.40
W. Africa - C.	-	-	-	-2.87*

Notes: One (*), two (**) and three (***) asterisks indicate significance at the 10%, 5%, and 1% levels. Critical values are: -2.58 (10%), -2.89 (5%), and -3.51 (1%) (Fuller, 1976).

Concluding Remarks

This paper examined the degree to which price linkages within the world market of cotton have improved over the last decade. Weekly data from August 15, 1985 to December 24, 1987 (122 observations) and August 3, 1995 to January 9, 1997 (73 observations) from US, Greece, Central Asia, and West Africa were utilized.

According to the goodness of fit criterion (i.e. the R^2) in almost all cases a substantial improvement in price linkages has taken place. For example, while on average, about 50% of the variability of price in one region was explained by the variability of another region's price in period 1, in period 2 the variability explained increased to 75%. Furthermore, the main source of this improvement appears to be a result of short-run price transmission rather than long-run comovement.

TABLE II: The Goodness of Fit

	A Index	*US*	*Greece*	*W. Africa*	*C. Asia*
Period 1					
A Index	–	0.49	0.40	0.80	0.88
US	0.50	–	0.18	0.32	0.43
Greece	0.46	0.16	–	0.44	0.44
W.	0.80	0.31	0.36	–	0.72
C. Asia	0.88	0.42	0.35	0.71	–
Period 2					
A Index	–	0.73	0.84	0.81	0.87
US	0.74	–	0.62	0.73	0.76
Greece	0.85	0.59	–	0.62	0.80
W.	0.81	0.72	0.62	–	0.75
C. Asia	0.87	0.75	0.79	0.74	–

Notes: Goodness of fit is the R^2 of equation (4).

TABLE III: The Short-run Effect

	A Index	*US*	*Greec e*	*W. Africa*	*C. Asia*
Parameter estimate of β_1 in period 1					
A Index	–	0.51	0.54	0.94	0.83
US	0.98	–	0.49	0.81	0.80
Greece	0.70	0.37	–	0.67	0.56
W.	0.85	0.39	0.49	–	0.72
C. Asia	1.05	0.53	0.58	1.00	–
Parameter estimate of β_1 in period 2					
A Index	–	0.49	0.85	1.05	0.93
US	1.48	–	1.22	1.68	1.50
Greece	1.00	0.49	–	0.99	0.97
W.	0.78	0.41	0.63	–	0.72
C. Asia	0.94	0.49	0.82	1.02	–
Test of equality of β_1 between the two periods: p-					
A Index	–	0.90	0.00	0.16	0.17
US	0.00	–	0.00	0.00	0.00
Greece	0.00	0.39	–	0.02	0.00
W.	0.29	0.90	0.19	–	0.92
C. Asia	0.13	0.80	0.02	0.80	–

Notes: All reported coefficients are significant at the 1% level. *p*-value is the significance level of the *F*-statistic of the hypothesis that β_1 in (4) is the same in the two periods.

TABLE IV: Dynamic Adjustment

	A Index	*US*	*Greec e*	*W. Africa*	*C. Asia*
Parameter estimate of $(1 - \beta_3)$ in period 1					
A	–	0.01	0.01		
US	0.02	–		0.03^{*}	0.02
Greece		0.02	–		
W.		0.00	0.00	–	
C. Asia		0.00	0.02		–
Parameter estimate of $(1 - \beta_3)$ in period 2					
A	–	0.02			
US	0.10^{**}	–			
Greece		0.04^{*}	–		
W.		0.00		–	
C. Asia	0.11^{**}	0.02	0.08^{*}		–
Test of equality of $(1 - \beta_3)$ between the two periods:					
A	–	0.60	0.02	0.71	0.96
US	0.12	–	0.15	0.21	0.14
Greece	0.22	0.52	–	0.99	0.51
W.	0.34	0.81	0.06	–	0.91
C. Asia	0.61	0.51	0.44	0.92	–

Notes: *p*-value is the significance level of the *F*-statistic of the hypothesis that $(1 - \beta_3)$ in (4) is the same in the two periods.

TABLE V: Number of Periods Required to Achieve 95% of the Adjustment

	A Index	*US*	*Gree ce*	*W. Africa*	*C. Asia*
Estimate of n in period 1					
A	–	n.c.	n.c.	0.8	8.7
US	n.c.	–	54.8	46.4	n.c.
Greece	12.3	n.c.	–	11.2	11.4
W.	4.6	n.c.	n.c.	–	9.6
C. Asia	0.4	n.c.	n.c.	0.0	–
Estimate of n in period 2					
A	–	n.c.	4.5	0.0	2.9
US	22.3	–	11.5	26.3	27.0
Greece	0.0	54.3	–	0.0	0.0
W.	9.6	n.c.	17.2	–	9.0
C. Asia	1.5	n.c.	16.2	0.0	–
Test of equality of n between the two periods: p-					
A	–	n.a.	0.00	0.32	0.37
US	0.00	–	0.00	0.00	0.00
Greece	0.01	0.55	–	0.06	0.00
W.	0.39	n.a.	0.13	–	0.99
C. Asia	0.30	n.a.	0.07	0.95	–

Notes: *p*-value is the significance level of the *F*-statistic of the hypothesis that $(1 - \beta_3)$ and β_1 in (4) (and hence k and n) is the same in the two periods. n is calculated as $[log(0.05) - log(1-\beta_1)]/ log\beta_3$ – a result of setting $k = 0.95$ and solving (9) for n. 'n.c.' indicates that long-run convergence never takes place as the error-correction parameter is not significantly different from zero. 'n.a.' indicates that the test is not reported because the respective prices did not converge..

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