International Steam Coal Market Integration

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Abstract

This paper examines the hypothesis that there is a single economic market for the international steam coal industry and investigates the degree of steam coal market integration over time. The long-run relations between international steam coal prices are tested through cointegration analysis and the Kalman Filter analysis is employed to examine the convergence path of the price series. A regression test of convergence is employed to test for group convergence within a panel of steam coal exporting countries. Monthly F.O.B. prices for Australia, China, Colombia, Indonesia, Poland and South Africa between January 1995 and July 2007 are considered. Using three different tests, we obtained a consistent conclusion – the international steam coal market is generally integrated.

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1. INTRODUCTION

In this paper, two related questions are addressed. Various authors have suggested, explicitly or implicitly, the existence of a unified global economic market for coal (Ellerman, 1995; IEA, 1997; Humphreys and Welham, 2000). Are these claims supported by empirical evidence? This is the first question that this paper will address. Having given a proper definition to “economic market”, how integrated, or unified, are the coal markets in different geographic regions? This is the second question that we will address. While the first question yields a yes or no answer, the second question concerns the strength of integration among the markets. We will investigate these questions by examining the relations between the coal export prices for Australia, China, Colombia, Indonesia, Poland and South Africa.

As coal is bulky and costly to move around, a logical and possible outcome is that there should be a few geographically separated markets, in particular, Europe, Asia and America. Differentials in transportation costs will make it more difficult for some suppliers to reach certain markets. However, there are several links between these regional markets. According to Ellerman (1995) and IEA (1997), the link is the US. For Humphreys and Welham (2000), the link is South Africa. Warell (2006) includes Australia as one of the links as well. It is true that imports and exports will normally flow between nearby countries, but the important message common to these studies is that due to the development of seaborne coal trade, suppliers can ship the coal almost anywhere on the globe, as long as there is enough demand and the price is right. Should these links be strong enough, it is reasonable to think about the market for coal as a global one, which is integrated or unified.

A number of studies have examined the degree of market integration in the US, European and global natural gas markets (King and Cuc, 1996; Serletis, 1997; Serletis and Herbert, 1999;

Asche, Osmundsen and Tveteras, 2002 and Siliverstovs, L'Hegaret, Neumann and von Hirschhausen, 2005). Ripple (2001) analyses the degree of market integration for the US West Coast with the US Gulf Coast and Asia. Expanding the horizon, Bachmeier and Griffin (2006) evaluate the degree of market integration both within and between crude oil, coal and natural gas markets. Examining a similar question to our paper, Warell (2006) tests the hypothesis of a single international market for coal.

Except for King and Cuc (1996), all of the above studies relied heavily on the use of cointegration techniques. Depending on the type of cointegration tests employed and the way the empirical models were defined, the studies obtained the most appropriate results to make inferences to the research questions. The basic methodology is that two or more markets can be concluded as integrated when cointegration is found among the price series in question. King and Cuc (1996) is an exception whereby the authors focused their study on price convergence rather than cointegration.[1]

2. METHODOLOGY

2.1 The Delineation of Markets

The definition of a market has received little attention from economists historically. As noted by Stigler (1982), except for some casual examination with cross elasticities of demand and supply, the determination of markets was an undeveloped area at both the theoretical and empirical level. The situation has changed since the early 1980s. Let us consider three common methods in market determination.

[1] According to King and Cuc (1996), convergence means that "the difference between two or more series should become arbitrarily small, or converge on some constant, with the passage of time".

The definition of market was traditionally based on the cross price elasticity of demand that measures the responsiveness of the change in demand for a product to changes in the price of another product. If the cross price elasticities are high, the products are considered to be part of the same market. Yet, an immediate problem arises – how high should cross elasticity be for us to conclude that the products belong to the same market? Moreover, Werden (1998) points out that when there are many product brands, cross elasticities between any pair of product may be small. When no individual brand has any significant market power, a small increase in the price of one brand may lead to considerable substitutions to the others, with each brand gaining too small a fraction for the cross elasticity measures to conclude that any of them are in the same market. Werden (1998) argues, rather than looking at the cross elasticities of product pairs, one should actually consider the collective competitive significance of all substitutes.

Another commonly used method to define a market area is product flows. Elzinga and Hogarty (1973) suggest defining markets basing on shipments data in physical terms. The Elzinga-Hogarty test consists of the little in from outside (LIFO) test and the little out from inside (LOFI) test. If 75% or more of the total sales in the hypothetical market area are shipped from plants within the area and 75% or more of the shipments by firms in this area go to customers in the area, the test will conclude that the hypothetical area constitutes an economic market. However, Kaserman and Zeiel (1996) argue that the critical values (75%) are not justified either theoretically or practically. Another important criticism towards the Elzinga-Hogarty test is that it does not account for potential competition. As mentioned by Kaserman and Zeiel (1996), the potential shipments from outside may already be sufficient to temper the pricing decisions of producers inside the area, making physical shipments unnecessary. In the case where there are no physical shipments between two areas but the cross-elasticity of demand is high, the two areas do belong to the same economic market. Yet, the Elzinga-Hogarty test will likely to conclude otherwise, providing a misleading result.

Stigler and Sherwin (1985) point out that cross elasticity tests involve additional complexity and stringent data requirements. They also show that trade flows between two areas cannot be used to determine whether or not the areas constitute a market. They favour the use of price correlations to define product markets. The rationale behind is that prices for products in the same market cannot move too much out of line with one another – a prediction that logically follows the Law of One Price.

2.2 A Price-based Definition

Throughout this paper, the price-based definition of market is adopted. Stigler (1990) states that, "a market, according to the masters, is the area within which the price of a commodity tends to uniformity, allowance being made for transportation costs". By using this market definition, we identify participants in the market that establish the price for the product in the market area. It must be noted that we cannot use this market definition to determine the underlying market structure. Without further information, we cannot tell whether the different areas are integrated because of the competitive market behaviour, the formation of spatial monopoly or oligopoly, or the changes in supply conditions of the market areas.

In competitive markets with zero transportation and transaction costs, no trade barriers and each market is connected to every other market, the Law of One Price states that a single price will hold in all market locations for identical products. Officer (1986) notes that a sufficient condition for the Law of One Price to hold is that the market is perfectly competitive, since this assures the existence of perfect arbitrage. In the real world, however, transaction costs (including search and information costs, bargaining costs and policing and enforcement costs) are usually not zero and markets are usually not perfectly competitive. The existence of these divergences will lead to deviations from prediction of the Law. Market power in the form of monopolistic or oligopolistic practices may weaken the correlation between the prices of similar goods. This is because a monopolist can practice price discrimination and charge different prices for buyers with different demand elasticities. For non-colluding oligopolists, especially at times of high freight costs, there will be incentives to absorb part of the freight costs and charge more distant buyers a lower F.O.B. price (making the C.I.F. price more competitive). Product differentiation may reduce the substitutability of the product for the consumers. If the commodities in question are not identical, the elasticity of substitution between these commodities will need to be sufficiently high for the Law of One Price to hold. In relation to steam coal, this may be seen when utilities must differentiate between coals of differing qualities as inputs (for example, heat content or moisture content), so the substitutability of the coals is low. As a result, the price relation between these coals is expected to be low. Yet, as technology advances, utilities are more and more able to accept coal of differing quality. Hence, product differentiation may not seriously affect the Law of One Price implications in the steam coal industry.

On the other hand, assuming that all firms are Cournot competitors, Ohta (1988) shows that under the spatial oligopoly setting with market overlap, the prices of a homogenous commodity will tend to uniformity in the overlapping market area. The price in the overlap area is a function of average marginal costs of production and transportation of all producers serving that area. It follows that for a market point that is on average more distant from the producers will have to pay a higher price – a prediction that is consistent with that of the Law of One Price.

Stigler and Sherwin (1985) demonstrate that in the situation where the transportation cost is so substantial that it separates two (originally) competitive markets, when either market is monopolised, the local price can be raised to the point where it equals the offshore competitive price plus transportation costs. According to the price definition, these markets are now integrated. Massey (2000) also remarks that a profit-maximising monopolist will generally set its price to the point that other products (or products from another area) become close substitutes.

The phenomenon of price uniformity is consistent with many different market structures. The key issue in our definition of market is whether or not a consumer (coal buyer) can purchase the same good (coal) for the same price from different sellers in the area under consideration. If the products are not identical, the question becomes whether or not the price differentials of the products are stable over time.

The price test can be formulated as a stochastic equation as follows:

$$P_{j,t} = \alpha_{ij,\,t} + \beta_{ij,t} P_{i,t} + \varepsilon_{jt} \qquad (1)$$

where $P_{i,t}$ is the natural log of price in market i in period t, $P_{j,t}$ is the natural log of price in market j in period t, $\alpha_{ij,\,t}$ captures transaction costs and quality differences, $\beta_{ij,t}$ represents the degree of market integration and ε_t is a white noise. When $\beta_{ij,t} = 0$, there is no relation between the two markets. When $\beta_{ij,t} = 1$, the prices are proportional and the relative price is stationary (i.e. the price differential is stationary). In general, the closer $\beta_{ij,t}$ is to 1, the more integrated are the two markets at time t.

2.2 Convergence and Cointegration

In the empirical literature, the cointegration technique is a natural choice in testing for market integration. When the price series are cointegrated, it is often concluded that the markets are integrated because there exists a stable long-run relation between the prices. On the other hand, if one fails to detect cointegration, it will be concluded that the markets are not integrated. In this paper, the Johansen (1990) test will be used to detect any cointegrating relation among the price series. Stability of the cointegration relations will be assessed through recursive analyses.

Some previous studies have pointed out that cointegration relies on an implicit assumption that the structural relation among variables is fixed over the time period in question.[2] In other words, it is most useful for detecting already integrated markets. When the prices in question are converging, the market is progressing towards integration. Yet, conventional cointegration tests will probably reject cointegration, even though convergence has occurred by the end of the sample period. This means that the cointegration technique cannot be used when there are dynamic structural changes in the market nor can it detect market integration when the integrated period is not a long enough subset of the period under study. One can try to break down the data into different sub-samples and repeat the cointegration test, but the degrees of freedom will quickly be exhausted and the test will lack power.

2.4 Kleit's Critique

Kleit (2001) criticises the use of price correlation and cointegration analysis in assessing the degree of market integration. He argues that the cointegration test simply presents a "Yes or No" answer – the price series are either cointegrated or not. Kleit's critique is valid for the papers that were cited in his paper, but not entirely correct in general. One can actually go much further and assess "how integrated are the markets" by testing if the long run coefficients (β in our case) are close or statistically equal to unity. This can be done, for example, in the Johansen (1990) procedure by using a likelihood ratio test. Also, as we will discuss shortly in the next sub-section, one can examine the evolution path of market integration by the use of Kalman filter. These techniques can all enable the researcher to evaluate both the qualitative and quantitative side of market integration.

[2] These studies include: Hall, Robertson and Wickens (1992), Caporale and Pittis (1993) and King and Cuc (1996).

2.3 The Kalman Filter

In order to obtain a better picture of the degree of market integration over time, we take into consideration the possibility of dynamic structural change in the series. Employing a similar approach as King and Cuc (1996), the Kalman filter is used to complement the cointegration results. By using the Kalman Filter, we do not have to break down the data into sub-samples while we can still examine the dynamic behaviour of the entire dataset.

The Kalman Filter uses temporal series of observable variables ($P_{j,t}$ and $P_{i,t}$) to compute the optimal estimates of $\alpha_{ij,\,t}$ and $\beta_{ij,t}$ in equation (1) for each time period. Following Harvey (1993), let us first define the state vector as $\pi_t = [\ \alpha_{ij,t}\ \beta_{ij,t}\ ]$ and a vector $Z_t = [\ 1\ P_{i,t}\ ]$. Then the *measurement equation* is formulated as

$$P_{j,t} = \pi_t Z_t' + \varepsilon_t \qquad (2)$$

where $E(\varepsilon_t) = 0$ and $Var(\varepsilon_t) = H_t$. The measurement equation represents the relation between price series and the state variables. The corresponding *transition equation* is:

$$\pi_t = \pi_{t-1} + \eta_t \qquad (3)$$

where $E(\eta_t) = 0$ and $Var(\eta_t) = Q_t$. This equation describes the dynamics of the state variables. The error terms are assumed to be uncorrelated with each other and the initial state for all time periods.

Although π_t is not directly observable, it can be determined from past values of π_t. First, the transition equation is used to compute $\pi_{t/t-1}$, estimates of π at time t using information up to time t-1. The *prediction equations* used in the process are:

$$\pi_{t/t-1}^{\ *} = \pi_{t-1}^{\ *} \qquad (4)$$

and

$V_{t/t-1} = V_{t-1} + Q_t$ (5)

where $Vt_{-1} = E[(\pi_{t-1} - \pi_{t-1}^{*})(\pi_{t-1} - \pi_{t-1}^{*})']$ and asterisks signifying the optimal predictor for the parameter. The predicted values are then substituted into the measurement equation to compute $P_{j,t}$. Next, when information at time period t becomes available, the estimate of π_t and V_t are updated using the *updating equations*:

$\pi_t^{*} = \pi_{t/t-1}^{*} + V_{t/t-1}Z_t' (Z_tV_{t/t-1}Z_t' + H_t)^{-1} (P_{jt} - Z_t'\pi_{t/t-1}^{*})$ (6)

and

$V_t = V_{t/t-1} - V_{t/t-1}Z_t' (Z_tV_{t/t-1}Z_t' + H_t)^{-1} (Z_tV_{t/t-1})$ (7)

This process is performed recursively until all the information in the dataset is exhausted.

2.4 A Regression t-Test of Convergence

Phillips and Sul (2007) propose a mechanism for modelling and analysing economic transition behaviour in the presence of common growth characteristics. The formulation is useful in measuring transition towards a long-run growth path or individual transitions over time relative to some common trend, representative or aggregate variable. The test does not rely on any assumptions of trend stationarity or stochastic nonstationarity in the data. Moreover, the formulation of the test allows for transitional divergence. We follow their approach to test for convergence of the international steam coal prices. Letting the price series for country *i* be P_{it}, the cross sectional variance ratio C_1/C_t is constructed, where

$$C_t = \frac{1}{N}\sum_{i=1}^{N}(c_{it}-1)^2 \quad \text{and} \quad c_{it} = \frac{P_{it}}{N^{-1}\sum_{i=1}^{N}P_{it}}.$$

Next, the following regression is estimated and an autocorrelation and heteroskedasticity robust one-sided t-test is applied on $\hat{b}$.

$$\log\left(\frac{C_1}{C_t}\right) - 2\log L(t) = \hat{a} + \hat{b}\log t + \hat{u}_t \qquad (8)$$

where $L(t) = \log(t+1)$ and $t = [kT], [kT]+1, \ldots, T$ with $k > 0$. k represents the fraction of data that is excluded in the regression stage of the test. It helps to focus attention in the test on what happens as the sample size gets larger. $L(t)$ is a slowing varying function for which $L(t) \to \infty$ as $t \to \infty$. Other possible choices for $L(t)$ include $\log(t+1)^2$ and $\log(\log(t+1))$. As shown by Phillips and Sul (2007), $k = 0.3$ is a satisfactory choice in terms of both size and power. $\hat{b}$ provides an estimate for the individual specific decay rate θ where $\hat{b} = 2\hat{\theta}$. If the price series in the panel are converging, $c_{it} \to 1$ and $C_t \to 0$ as $t \to \infty$. Accordingly, $\log(C_1/C_t)$ diverges to ∞, either as $2\log L(t)$ when $\theta = 0$ or as $2\theta \log t$ when $\theta > 0$. As a result, the null of convergence can be tested in terms of $\theta \geq 0$ and since θ is a scalar, the null can be tested using a simple one-sided t-test. The null of convergence is rejected if $t_{\hat{b}}$ is less than the critical value.

3. EMPIRICAL ANALYSIS

3.1 Data

For the empirical analysis, thermal coal F.O.B. prices for Australia (Newcastle/Port Kembla), China (Qinhuangdao), Colombia (Bolivar), Indonesia (Kalimantan), Poland (Baltic Ports) and South Africa (Richards Bay) are used. The prices are quoted in US dollars per metric tonne. Observations are monthly and the dataset covers the period from 1995:1 to 2007:7 (151 observations). These countries are included because they are the major exporters of steam coal, and reliable price data are available. An advantage of using F.O.B. prices is that the information contained in the prices is more "pure" because it reflects only the price for coal from an origin (allowing for land transportation costs from mines to the export port). The final delivered prices will differ when the coal is shipped to different destinations due to different transportation and transaction costs. The use of F.O.B. prices avoids the complications brought by differing transportation costs, making our analysis more consistent with the Law of One Price, since the law is about prices *after allowance being made to transportation costs*.[3]

Figure 1. Thermal Coal Prices

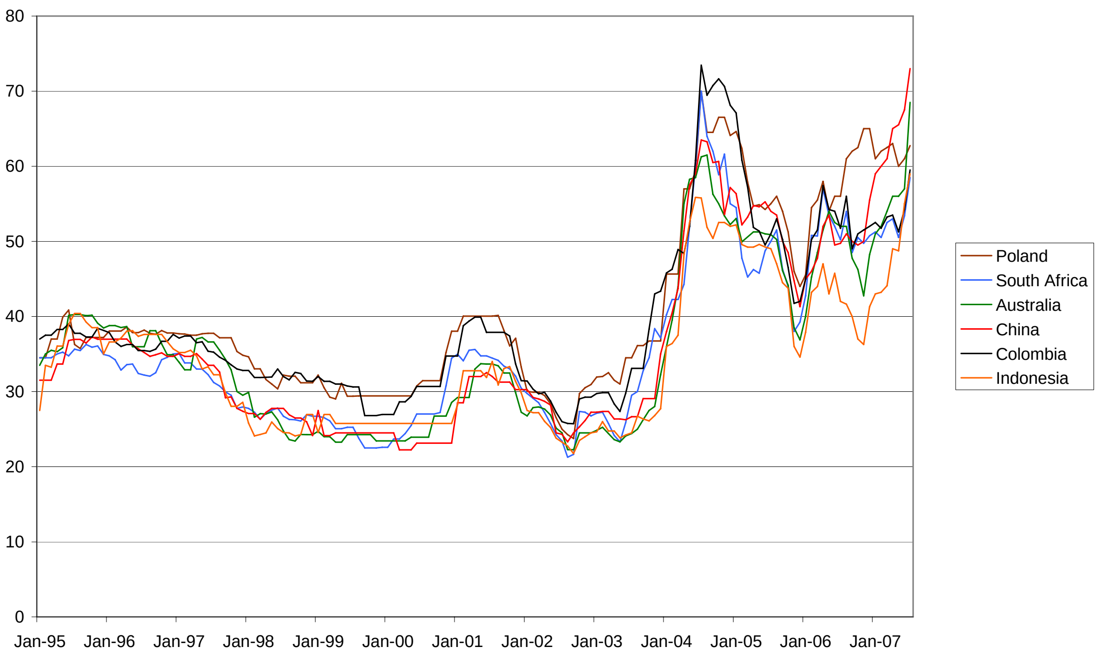


Figure 1 shows the F.O.B. thermal coal prices in levels. It can be observed that in general the prices exhibit very similar movement patterns throughout the sample. The prices fluctuated between $20 and $40 from 1994 to 2004. The prices spiked in mid-2004, declined until early 2006, when they started surging again.

[3] An implicit assumption in this analysis is that there is no freight absorption by coal exporters, where exporters charge a lower F.O.B. price for more distant buyers. Freight absorption can be a result of strong bargaining power from the buyer's side or spatial price discrimination on the seller's side.

3.2 Unit Root Tests

To ensure that our cointegration tests are meaningful, we have to confirm that the price series are integrated of order one prior to testing for cointegration. In both the raw and log-transformed data, it is found that all the price series have a non-zero mean. On the other hand, it is not entirely clear whether or not they exhibit trends. Adding an unnecessary constant or trend term to a model will affect the power of the test. Consequently, a sequential testing procedure is adopted.[4] The lag lengths for the augmented Dickey-Fuller (ADF) tests are chosen so that the error terms are serially uncorrelated (according to the Lagrange Multiplier and Ljung-Box autocorrelation tests) and the AIC and BIC information criteria are minimized. For the Phillips-Perron (PP) test, the lag lengths are chosen according to the formula $12(T/100)^{1/4}$, where T is the number of observations. The PP test uses a non-parametric correction to the t-test statistics to account for autocorrelation in the regression model. All the variables enter the regression in natural log form.

The test results are reported in Table 1. The null hypothesis is that a unit root exists, while the alternative hypothesis is that the series is stationary. According to both the ADF and PP tests, we can see that all the price series are non-stationary in level but stationary in first-difference. It means that all of the series are integrated of order one.

[4] The unit root regression model is: $\Delta X_t = \alpha_0 + \alpha_1 X_{t-1} + \alpha_2 t + \gamma_1 \Delta X_{t-1} + \ldots + \gamma_{p-1} \Delta X_{t-p-1} + e_t$.
First, we test for $H_0: \alpha_1 = \alpha_2 = 0$ using the Φ_3 statistic. Φ_3 has a non-standard F-distribution. We then test for $H_0: \alpha_1 = 0$ using the τ_t statistics. The τ_t statistics is compared to a standard t-critical value if Φ_3 is significant and to a non-standard t-critical value if Φ_3 is insignificant. Second, if τ_t is not rejected, we will proceed to test for $H_0: \alpha_1 = \alpha_2 = 0$ using the Φ_1 statistic. If the null is rejected, compare τ_μ to a standard t statistic; if the null is not rejected, compare τ_μ to a non-standard t statistic. Finally, if τ_μ is not rejected, compare τ to a non-standard t statistic.

Table 1. Augmented Dickey-Fuller and Phillips-Perron Tests

		Augmented Dickey-Fuller				
		Constant, Trend		Constant		No Const./ Trend
Variable	Lags	Φ_3	τ_t	Φ_1	τ_μ	τ
Level						
Australia	1	2.05	-1.32	0.44	-0.56	0.70
China	1	1.76	-0.86	0.71	0.16	1.20
Colombia	1	2.25	-1.85	0.65	-0.96	0.55
Indonesia	0	0.98	-1.03	0.72	-0.60	1.00
Poland	6	2.99	-2.35	1.18	-1.42	0.48
South Africa	3	2.48	-1.86	0.54	-0.82	0.57
First difference						
Australia	0	29.49***	-7.65***	28.09***	-7.49***	-7.47***
China	0	55.72***	-10.56***	53.70***	-10.36***	-10.28***
Colombia	0	43.90***	-9.31***	13.36***	-9.31***	-9.31***
Indonesia	0	58.90***	-10.83***	57.69***	-10.74***	-10.74***
Poland	5	6.88**	-3.71**	6.69***	-3.66***	-3.62***
South Africa	2	23.32***	-6.82***	22.51***	-6.61***	-6.69***

		Phillips-Perron				
		Constant, Trend		Constant		No Const./ Trend
Variable	Lags	$Z(\Phi_3)$	$Z(\tau_t)$	$Z(\Phi_1)$	$Z(\tau_\mu)$	$Z(\tau)$
Level						
Australia	2	1.65	-1.08	0.44	-0.39	0.82
China	2	1.74	-0.91	0.69	0.11	1.17
Colombia	2	1.98	-1.71	0.59	-0.87	0.60
Indonesia	2	1.19	-1.29	0.82	-0.90	0.85
Poland	2	1.59	-1.56	0.64	-0.72	0.82
South Africa	2	2.26	-1.79	0.57	-0.86	0.57
First difference						
Australia	2	29.06***	-7.60***	27.63***	-7.43***	-7.40***
China	2	56.71***	-10.64***	54.71***	-10.46***	-10.38***
Colombia	2	44.15***	-9.39***	43.61***	-9.34***	-9.34***
Indonesia	2	59.06***	-10.85***	57.89***	-10.76***	-10.76***
Poland	2	61.08***	-11.05***	60.88***	-11.04***	-11.01***
South Africa	2	44.19***	-9.39***	43.53***	-9.33***	-9.33***

From MacKinnon (1991), critical values at 5% significance with 140 observations are -3.44 (τ_t), -2.88 (τ_μ) and -1.94(τ). The Dickey-Fuller (1979) critical values at 5% with 100 observations are 6.49 (Φ_3) and 4.71 (Φ_1). ** and *** indicate that the value is significant at 5% and 1% significance level, respectively.

3.3 Cointegration

Because of its geographic location, South Africa is often regarded as the link between the Pacific and Atlantic regions. Hence, we take South Africa as an anchor and test for cointegration of each country against it. The Johansen (1990) procedure is adopted to detect cointegration in our dataset. The lag length is chosen to ensure that the error terms are independently and identically distributed, while the model is the most parsimonious. Structural indicator dummy variables are included (as necessary) in the models to account for unusually large fluctuations in the dataset to deal with non-normality of the residuals. In terms of the deterministic components of the model, unrestricted constants are restricted to the cointegrating relations. It is found that the estimation results are robust against different lags and deterministic structures.

The bivariate cointegration results are presented in Table 2. The null of no cointegrating vector ($r = 0$) is rejected for Australia, China, Colombia and Poland at 5% significance but not for Indonesia even at 10% significance. The null of one cointegrating vector ($r = 1$) cannot be rejected at 10% significance for all models. This indicates that there is one cointegrating vector between each of these price series (except Indonesia) and South Africa.

Table 2. Cointegration Test against South Africa

Country	Lags	p-r	r	Eig. Value	Trace	P-value
Australia	7	2	0	0.101	16.461	0.034
		1	1	0.008	1.193	0.275
China	5	2	0	0.133	21.619	0.004
		1	1	0.005	0.760	0.383
Colombia	8	2	0	0.108	16.912	0.029
		1	1	0.004	0.532	0.466
Indonesia		2	0	0.076	11.553	0.183
		1	1	0.001	0.144	0.704
Poland	4	2	0	0.134	21.228	0.005
		1	1	0.000	0.019	0.892

P-values are based on critical values from Osterwald-Lenum (1992).

The normalized long-run coefficients for the error-correction model are shown in Table 3. Linear restrictions on the cointegration coefficients are tested using likelihood ratio tests. The results are also reported in Table 4. The cointegration coefficients for Australia and Poland are not statistically different from unity. This means that these prices move proportionally with the South African series, and the relative prices are constant in the long run. The price differentials are stable, and they reflect only quality differentials and perhaps (constant) transaction costs brought by trade frictions. In other words, these markets are fully integrated. This provides strong evidential support for the market integration hypothesis. Although the cointegration coefficients for China and Colombia are significantly different from unity, the fact that these prices are cointegrated with South Africa still serves as an indication of market integration. Since, at the very least, these cointegrated prices share a common stochastic trend and they will not drift apart without limit in the long run.

Table 3. Normalized Long-run Coefficients [a]

Country	β	$H_0{:}\beta = 1$	P-value
Australia	1.113	$\chi^2 = 2.684$	0.101
China	1.174	$\chi^2 = 6.664$	0.010
Colombia	0.842	$\chi^2 = 10.706$	0.001
Poland	0.953	$\chi^2 = 1.073$	0.300

[a] Normalized such that South Africa is the explanatory (RHS) variable

3.4 Recursive Estimation

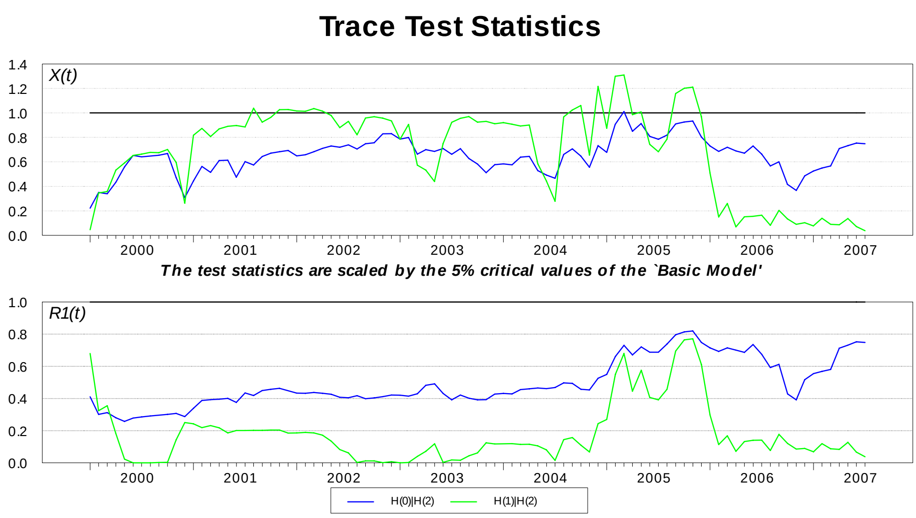
Recursive estimation is performed to test for the constancy of the long run coefficient, β, and the trace statistics in the models. Continuously adding observations to the base sample, the recursive calculations are done by re-estimating all parameters in each step (known as the X-form) and by re-estimating the long-run parameters only (known as the R1-form). Major differences between the results of the two forms may signify instability of the short run parameters. The base sample for the forward recursive estimation is from 1995:07 to 2000:01, while the sample for the β

("known beta") to be used for the β constancy tests is from 1995:07 to the end of the data set. When each sub-sample is taken individually, the test statistic is asymptotically distributed as χ^2 with (p-r)r degrees of freedom, where p is the number of variables and r is the cointegrating rank. The test statistics for both tests are scaled by the 5% critical value, such that a test statistic above unity indicates rejection at 5% significance. The results are graphed in Figure 3.

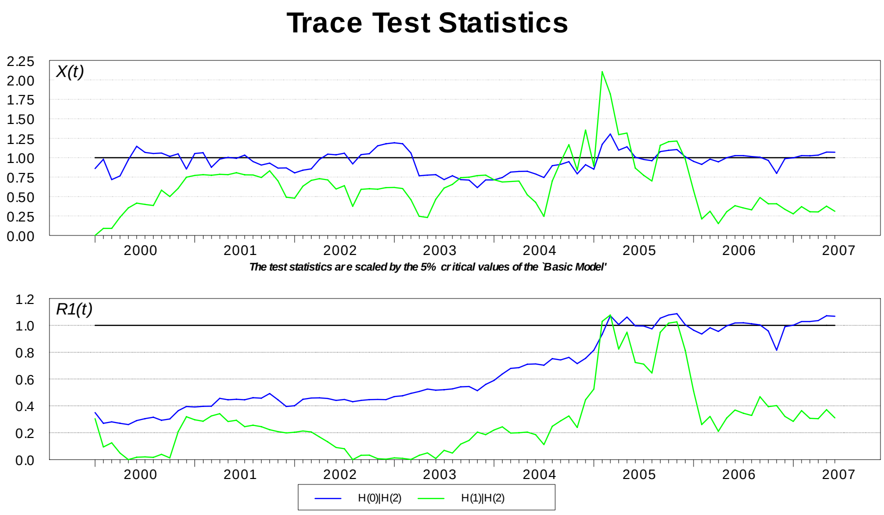
The recursive trace statistics indicate that the price series are cointegrated only after 2003 (Colombia) and 2004 (Australia, China and Poland). This holds for both the X-form and the R1-form estimations. Also, the Indonesian price is not cointegrated with the South African price over the whole time period considered. On the other hand, we can see that the long run coefficients for all the cointegration relations are stable throughout the whole sample in each step using the R1-form. However, if we consider all the parameters in the recursive estimation (X-form), we find fluctuations in the beginning of the recursion for Australia. However, starting from 2003, the long run coefficient stabilised. For Colombia and Poland, the long run coefficients are significantly different from the full sample estimate in the 2000-2002 period and for 2004.

Figure 3. Recursive Estimation

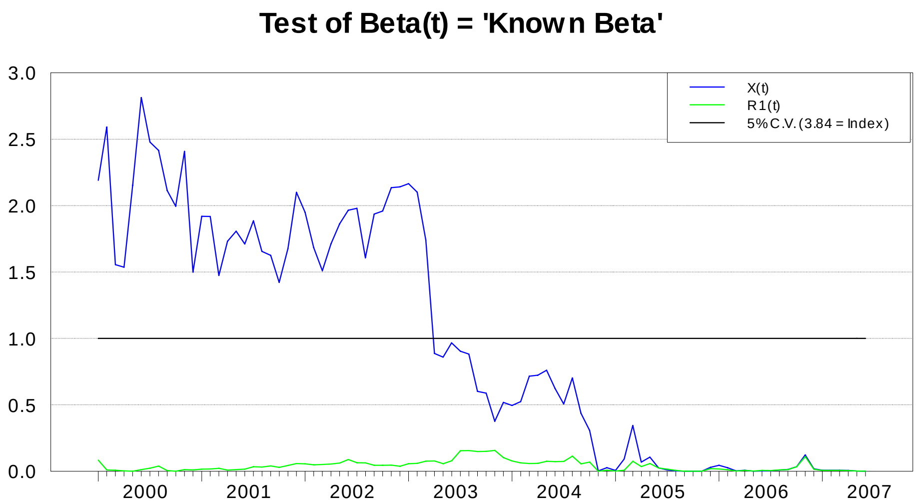
<u>Indonesia</u>



Australia

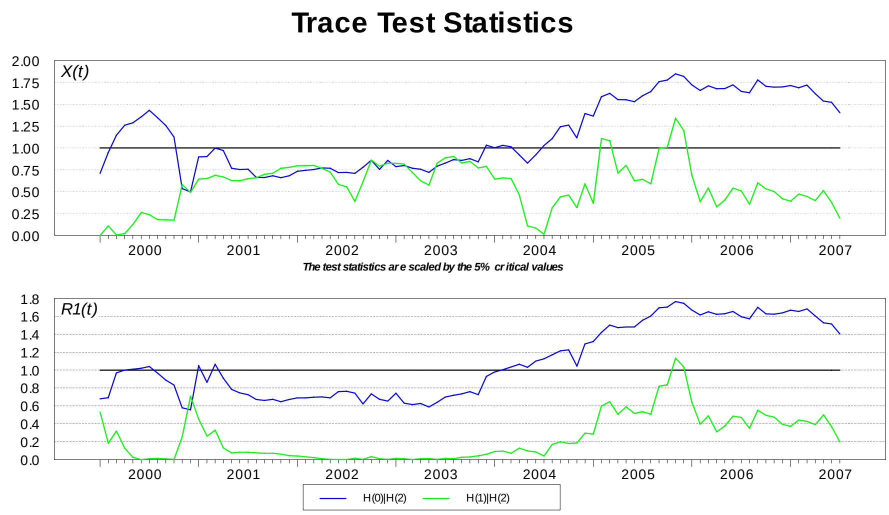


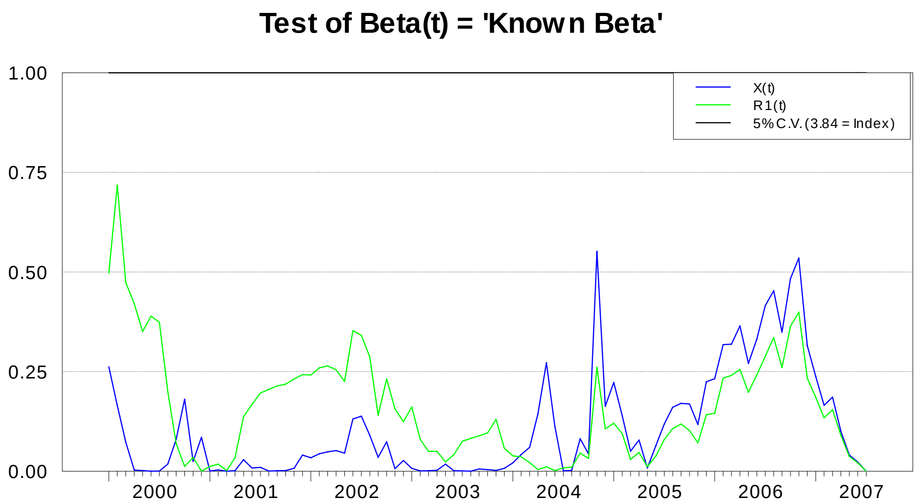
Trace Test Statistics
X(t)
2.25
2.00
1.75
1.50
1.25
1.00
0.75
0.50
0.25
0.00
2000
2001
2002
2003
2004
2005
2006
2007
The test statistics are scaled by the 5% critical values of the 'Basic Model'
R1(t)
1.2
1.0
0.8
0.6
0.4
0.2
0.0
H(0)|H(2)
H(1)|H(2)



Test of Beta(t) = 'Known Beta'
3.0
2.5
2.0
1.5
1.0
0.5
0.0
X(t)
R1(t)
5% C.V. (3.84 = Index)
2000
2001
2002
2003
2004
2005
2006
2007

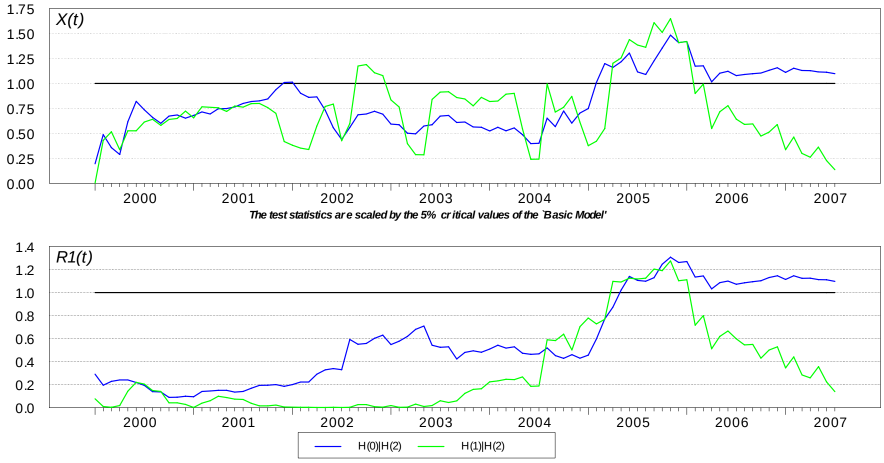
China


Trace Test Statistics
X(t)
2.00
1.75
1.50
1.25
1.00
0.75
0.50
0.25
0.00
2000
2001
2002
2003
2004
2005
2006
2007
The test statistics are scaled by the 5% critical values
R1(t)
1.8
1.6
1.4
1.2
1.0
0.8
0.6
0.4
0.2
0.0
2000
2001
2002
2003
2004
2005
2006
2007
H(0)|H(2)
H(1)|H(2)

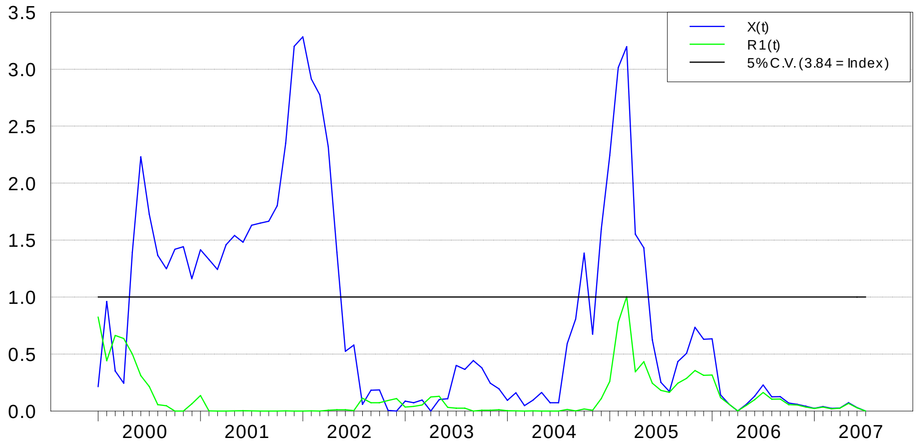

Test of Beta(t) = 'Known Beta'
1.00
0.75
0.50
0.25
0.00
X(t)
R1(t)
5% C.V. (3.84 = Index)
2000
2001
2002
2003
2004
2005
2006
2007

Colombia

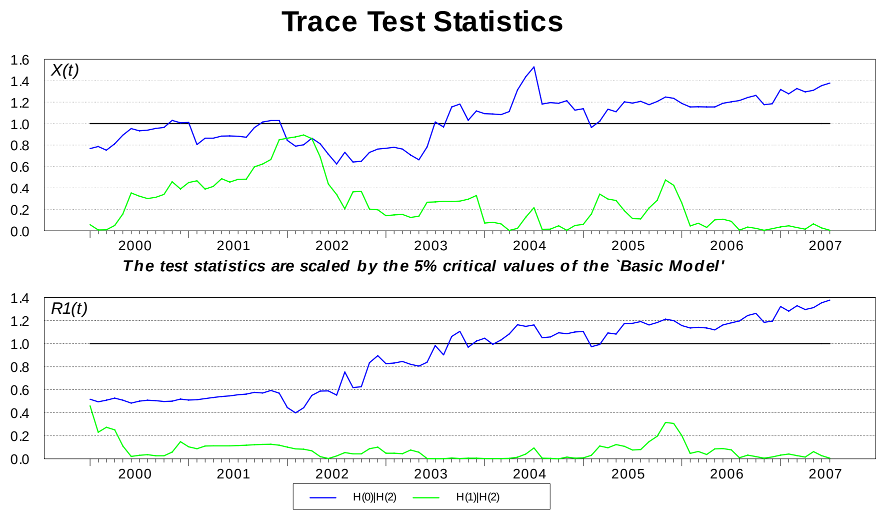
Trace Test Statistics



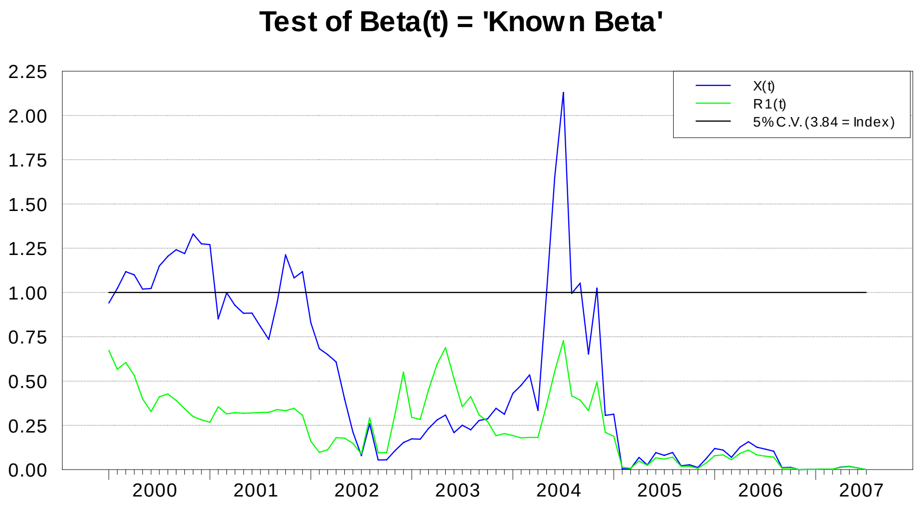
Test of Beta(t) = 'Known Beta'



Poland



Trace Test Statistics
X(t)
1.6
1.4
1.2
1.0
0.8
0.6
0.4
0.2
0.0
2000
2001
2002
2003
2004
2005
2006
2007
The test statistics are scaled by the 5% critical values of the `Basic Model'
R1(t)
1.4
1.2
1.0
0.8
0.6
0.4
0.2
0.0
2000
2001
2002
2003
2004
2005
2006
2007
H(0)|H(2)
H(1)|H(2)



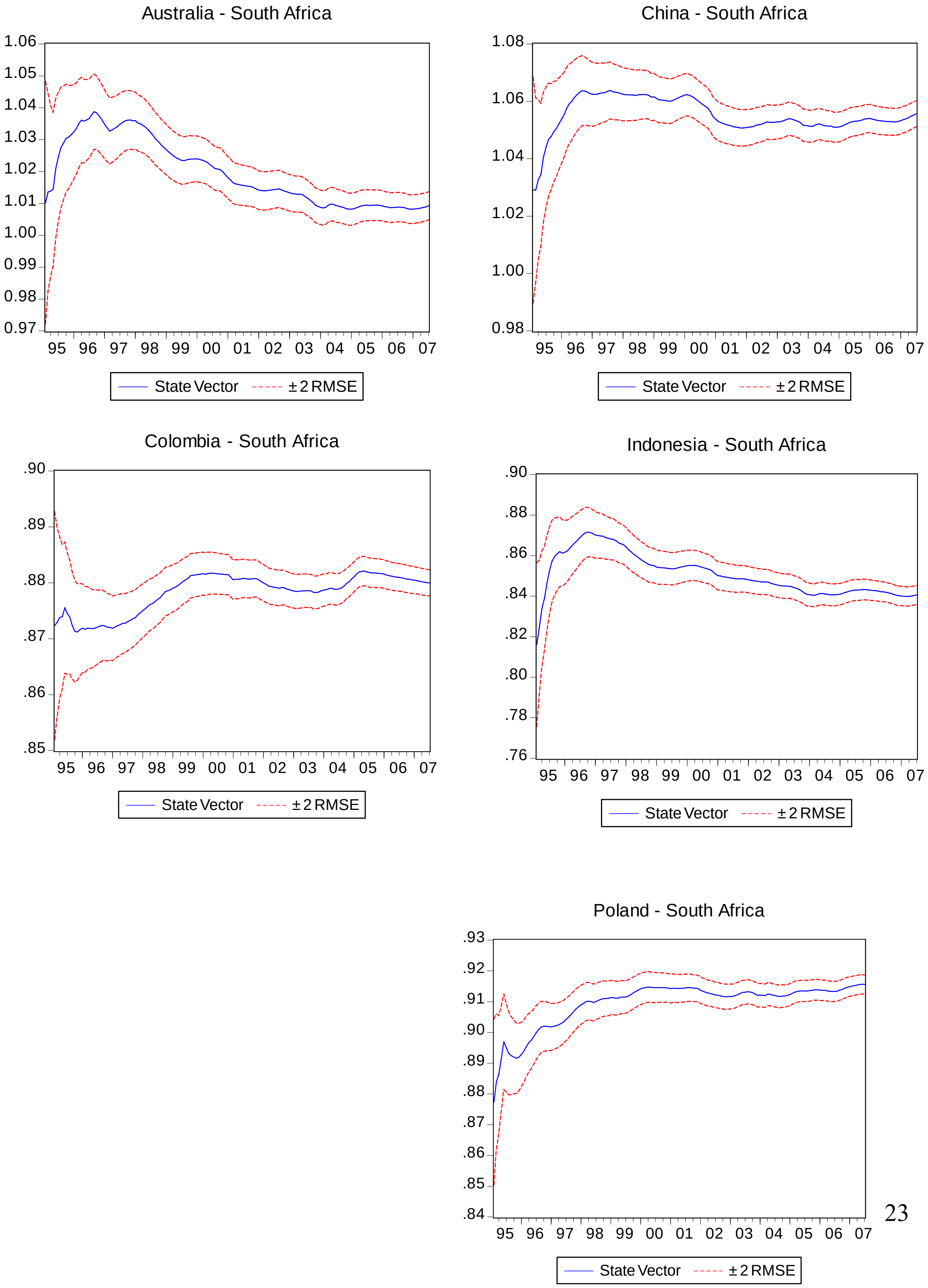
Test of Beta(t) = 'Known Beta'
2.25
2.00
1.75
1.50
1.25
1.00
0.75
0.50
0.25
0.00
2000
2001
2002
2003
2004
2005
2006
2007
X(t)
R1(t)
5% C.V. (3.84 = Index)

3.5 Kalman Filter Estimation

As mentioned in Section 2, the Kalman filter is used to complement the cointegration results by taking into account possible dynamic structural changes in the model. According to Bomhoff (1992), the Kalman filter allows for non-stationarity in the data. In a simulation study, Bomhoff (1992) finds that the coefficient estimates generated by the Kalman filter outperform the coefficient estimates generated by OLS. This conclusion holds true whether the non-stationary series are cointegrated or not. The Kalman filter in levels encompasses both cases and produces results that are useful when differencing the data will be appropriate. Therefore, data in levels will be used in our application of the Kalman filter. The initial conditions for the states and variances are set by diffuse prior. Following the method adopted by Koopman, Shephard and Doornik (1999), $\pi_{1|0} = 0$ and $V_{1|0} = \kappa I$, where κ is initially set to 10^6 and then adjusted for scale by multiplying by the largest diagonal element of the residual covariances. The time-varying paths of the β's are plotted in Figure 5.

Because the estimated time paths for β's may be sensitive to the choice of the initialization of the Kalman filter (especially for the earlier part of the sample), we consider the time paths starting from 1997. There are signs of price convergence for Australia and South Africa. The β coefficient moved towards unity rather steadily and stabilised at around 1.01 starting from 2004. This explains why in the recursive estimation, we found that the long run coefficient has stabilised from 2003 onwards. For China, β was fairly stable at 1.06 until it slightly dropped to 1.05 in 2000 and fluctuated around that level thereafter. The β coefficient for Colombia and Poland are also rather stable – they fluctuated around a mean of 0.88 and 0.91 respectively. In particular, β for Colombia peaked in 2000 and 2004 and then dropped back to a lower level. This is consistent with the recursive analysis. The time path of β for Indonesia reveals an extremely moderate rate of divergence from unity. The coefficient path stabilised in 2004. Finally, it must be noted that while there are signs of fluctuation in the integration coefficient, none of these changes are above 0.03 in absolute magnitude over the 10-year period.

Figure 5. Time-varying Paths of the Integration Coefficient (β)



5.6 Phillips-Sul Convergence Test

Finally, we consider the Phillips-Sul group convergence test. The hypotheses of the test are:

H_0: $\delta_i = \delta$ and $\theta \geq 0$
H_A: $\delta_i \neq \delta \; \forall \, i$ or $\theta < 0$

where δ_i measures the idiosyncratic distance between some common factor and the systematic part of the price series. The estimated equation for the log *t* regression as shown in equation (8) with $r = 0.3$ is

$$\log \frac{C_1}{C_t} - 2 \log t = -2.254 - 0.164 \log t$$
$$(-3.542) \quad (-1.124)$$

The test statistics are in parenthesis. At 10% and 5% significance, the critical values for the one-sided t-test are approximately -1.290 and -1.660, respectively. The null hypothesis of convergence is not rejected even at the 10% significance level. It implies that the cross sectional variance of the steam coal prices converges to zero as $t \to \infty$. This provides evidence of group convergence for international steam coal prices.

4. DISCUSSIONS AND CONCLUSIONS

The cointegration results show that the Australian, Chinese, Colombian and Polish steam coal prices are cointegrated with that of South Africa. The Kalman filter estimation further indicates that the price series exhibit a varying but high degree of integration. It means that buyers from all parts of the globe can make their purchase from these countries and expect the price differentials to reflect only differences in quality. Furthermore, the Phillips-Sul convergence test shows evidence of group convergence in the whole panel of countries we have chosen.

One seemingly conflicting result from these tests is that there is no strong evidence that Indonesia is integrated (or integrating) with the other steam coal exporters according to the cointegration and Kalman filter analyses. However, the Phillips-Sul test does show that Indonesia belongs to the same convergence group with all the other countries in our study. A possible reason for this "conflict" is that the formulation of the Phillips-Sul test allows for transitional divergence, i.e. the series can temporarily diverge from the general convergence path. If the divergence of Indonesia (if any) is just transitional, the test will not reject the convergence hypothesis. Also, the Phillips-Sul test considers convergence in terms of a group, instead of using a particular member as the anchor. Even though the Indonesian price is not cointegrated with that of South Africa, it does not necessarily mean that Indonesia is completely detached from the group. On balance, the international steam coal market can in general be regarded as integrated.

In a similar study, Warell (2006) tests the hypothesis of a unified economic market for coal using European and Japanese C.I.F. import prices. The study was separated for coking coal and steam coal. Warell (2006) finds that the steam coal prices for Japan and Europe are cointegrated for the sample period 1980:Q1 to 2000:Q3 and 1980:Q1 to 1989:Q4 but not in 1990:Q1 to 2000:Q3. The results suggest that the global coal market has become more regional in scope.

At first sight, our results appear to contradict Warell (2006) in the period 1990:Q4 to 2000:Q3. Yet, our study differs from Warell (2006) in at least two respects. First, the time period of our study only partly overlaps with that of Warell (2006). Our study can be viewed as an extension of Warell (2006). Keeping in mind that market integration and price convergence is an ongoing process, failure to detect cointegration does not necessarily imply that the market is not integrating. It is possible that the prices were actually converging but the level of price convergence or market integration is not strong enough for the cointegration test to pick up. Also, as indicated by the recursive trace statistics, cointegration occurred only in the later part of our sample. Taken together, the evidence suggests that the international steam coal market was not integrated until the early 2000's.

Second, our study uses F.O.B. prices of the major steam coal exporters rather than C.I.F. prices of Europe and Japan, as in Warell (2006). The information carried by C.I.F. prices used in Warell (2006) is quite different from F.O.B. prices, where an estimation of the transportation cost is included in the calculation of the C.I.F. prices. The quality of the price data depends crucially on the estimation of the freight rates. Also, the C.I.F. prices used by Warell (2006) are averages of coal with different quality, sold under different contractual arrangements and shipped from different sources. A great deal of noise is introduced into the price series through these averaging processes. Our use of F.O.B. prices avoids many of these complications. The use of F.O.B. prices is more consistent with the Law of One Price, since the law is concerned about the prices after netting out transportation costs.

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