Is there an East-West split in North-American natural gas markets?

Serletis, Apostolos

1997

IS THERE AN EAST-WEST SPLIT IN NORTH AMERICAN NATURAL GAS MARKETS?

This paper presents evidence concerning shared stochastic trends in North American natural gas (spot) markets, using monthly data for the period that natural gas has been traded on organized exchanges (from June, 1990 to January, 1996). In doing so, it uses the Engle and Granger (1987) approach for estimating bivariate cointegrating relationships as well as Johansen's (1988) maximum likelihood approach for estimating cointegrating relationships in multivariate vector autoregressive models. The results indicate that the east-west split does not exist.

INTRODUCTION

In the last decade, the North American natural gas industry has seen a dramatic transformation from a highly regulated industry to one which is more market-driven. The transition to a less regulated, more market-oriented environment has led to the emergence of different spot markets throughout North America. In particular, producing area spot markets have emerged in Alberta, British Columbia, Rocky Mountain, Anadarko, San Juan, Permian, South Texas, and Louisiana basins. Moreover, production sites, pipelines and storage services are more accessible today, thereby ensuring that changes in market demand and supply are reflected in prices on spot, futures, and swaps markets.

In a perfectly competitive industry the law of one price suggests that the difference in prices between any two markets should reflect the difference in transportation costs between the two markets. Because the natural gas molecule is identical when measured in terms of heating values, whether it comes from a well in Alberta or in the Gulf Coast, there is no reason that the law of one price should not apply to the natural gas industry. However, capacity constraints seem to be distorting North American natural gas markets in such a way that varying differentials emerge between spot prices, reflecting not only transportation costs but also supply and demand conditions in different areas.

Recently, King and Cuc (1996), in investigating the degree of natural gas spot price integration in North America, report evidence of an east-west split in North American natural gas markets. In particular, they argue that western prices tend to move together and similarly eastern prices tend to move together, but there seems to be a divergence between eastern and western prices. In other words, according to King and Cuc (1996) eastern and western prices seem to be determined by different fundamentals. King and Cuc (1996) use integration and (bivariate) cointegration analysis to measure natural gas price convergence, but mainly rely on a method of measuring convergence recently proposed by Haldane and Hall (1991). This method is based on the use of time-varying parameter (Kalman filter) analysis and is typically used to estimate regression type-models where the coefficients follow a random process over time.

In this study we investigate the robustness of the King and Cuc (1996) findings to alternative testing methodologies. In doing so, we test for shared stochastic price trends using current, state-of-the-art econometric methodology. In particular, we pay explicit attention to the time series properties of the variables and test for cointegration, using both the Engle and Granger (1987) approach as well as Johansen's (1988) (multivariate) maximum likelihood extension of the Engle and Granger approach. Looking ahead to the results, the tests indicate that the King and Cuc (1996) east-west split does not exist.

The paper is organized as follows. The second section provides some background regarding North American natural gas spot markets. The third section discusses the data and investigates the univariate time series properties of the variables, since meaningful cointegration tests critically depend on such properties. The fourth section tests for cointegration and presents the results. The last section concludes the paper.

THE NORTH AMERICAN NATURAL GAS SPOT MARKETS

The Alberta and British Columbia producing regions are part of the Western Canadian Sedimentary basin. In the case of Alberta, natural gas is transported from the field along the Nova Gas Transmission system for sale within Alberta as well as exported to eastern Canada and the United States. The British Columbia natural gas producing region is located mainly in northeastern British Columbia and natural gas is transported from the field along the Westcoast Gas Services system for sale in British Columbia and for export to the United States. Whereas gas exported from Alberta is resold in eastern markets in Canada and in the northeastern and midwestern United States, as well as in the western United States (specifically in California and the Pacific Northwest), British Columbia exports generally serve only markets in the Pacific Northwest and California.

The Rocky Mountain basin is a cluster of producing regions in the states of Wyoming, Utah, and Colorado. Pipelines in this area can transport production either east or west, although the eastward capacity has been constrained and thus, the majority of Rocky Mountain supplies are sold in the western markets in California and the Pacific Northwest. The San Juan basin is located in southwestern Colorado and northwestern New Mexico. Like the Rocky Mountain basin, eastward capacity is constrained which means that the majority of gas produced in this region is sold in western markets. The Rocky Mountain, San Juan, and Western Canadian Sedimentary basins comprise the western portion of the King and Cuc's (1996) east-west split (see Figure 1).

The Anadarko and Permian basins are located in the southwest United States (southern Oklahoma and western Texas, respectively). Supplies from both of these basins tend to be the marginal supplies of both eastern and western markets--that is, because natural gas from these basins can easily be shipped either east or west, it will be sold in the market with the highest price. The South Texas and Louisiana basins are located along the Gulf Coast. These areas are well served by a number of pipelines, which means that natural gas produced in these areas can reach most markets, either to the west or to the east. Price differentials, however, tend to support selling gas from these basins in eastern markets, such as the U.S. northeast and midwest. Although the Permian and Anadarko prices are set in both eastern and western markets, King and Cuc (1996) argue that these two basins with the South Texas and Louisiana basins comprise the eastern portion of the east- west split (see Figure 1).

In the next two sections, we investigate whether the price behavior of natural gas in different areas is similar. In particular, we use recent advances in the theory of nonstationary regressors to determine what trends in natural gas prices, if any, are common to Alberta, British Columbia, Rocky Mountain, San Juan, Anadarko, Louisiana, Permian and South Texas? Our definition of trend follows the cointegration literature. In particular, according to Beveridge and Nelson (1981) any time series characterized by a unit root can be decomposed into a random-walk and a stationary component, with the random-walk component being interpreted as the stochastic trend. Two series are said to share a trend if their stochastic trend components are proportional to each other. Clearly, a better understanding of the extent to which natural gas prices share trends might shed some light on the economic processes that determine natural gas prices.

THE DATA AND STOCHASTIC TRENDS

The data we use to test for shared stochastic natural gas price trends (from June 1990 to January 1996) are monthly bid-week prices reported by Brent Friedenberg Associates in the Canadian Natural Gas Focus. Bid week refers to the week during which pipeline nominations to transport gas take place. This is generally five days before the end of the month. Figure 2 shows the plots of natural gas prices in the western producing region of the King and Cuc (1996) east-west split--Alberta, British Columbia, Rocky Mountain, and San Juan basins. Figure 3 shows prices in the eastern producing region of the split--Anadarko, Louisiana, Permian, and South Texas basins.

The first step in testing for shared stochastic trends is to test for stochastic trends (unit roots) in the autoregressive representation of each individual time series. Nelson and Plosser (1982) argue that most macroeconomic and financial time series have a unit root (a stochastic trend), and describe this property as one of being "difference stationary" so that the first difference of a time series is stationary. An alternative "trend stationary" model has been found to be less appropriate. In what follows we test for unit roots using three alternative unit root testing procedures to deal with anomalies that arise when the data are not very informative about whether or not there is a unit root. In doing so, we choose to include only a constant (but not a time trend), since the series are not trending (see Figures 2 and 3).

The first three columns of panel A of Table 1 report p-values for the augmented Weighted Symmetric (ws) unit root test [see Pantula et al. (1994)], the augmented Dickey-Fuller (ADF) test [see Dickey and Fuller (1981)], and the Z(t[alpha]) nonparametric test of Phillips (1987) and Phillips and Perron (1988). These p-values are based on the response surface estimates given by MacKinnon (1994). For the WS and ADF tests, the optimal lag length was taken to be the order selected by the Akaike information criterion (AIC) plus 2-- see Pantula et al. (1994) for details regarding the advantages of this rule for choosing the number of augmenting lags. The Z(t[alpha]) test is done with the same Dickey-Fuller regression variables, using no augmenting lags. Based on the p-values for the WS, ADF, and Z(t[alpha]) test statistics reported in panel A of Table 1, the null hypothesis of a unit root in log levels cannot be rejected, except perhaps for the Permian price series.

To test the null hypothesis of a second unit root, in panel B of Table 1 we test the null hypothesis of a unit root [using the WS, ADF, and Z(t[alpha]) tests] in the first (logged) differences of the series. Clearly, all the series appear to be stationary in growth rates, since the null hypothesis of a unit root in the first (logged) differences of the series is rejected. We conclude that all the series are integrated of order one [or I(1) in the terminology of Engle and Granger (1987)].

TEST METHODS (AND CAPABILITIES) AND RESULTS

Since a stochastic trend has been confirmed for each price series, we now explore for shared stochastic price trends among these series by testing for cointegration (i.e., long-run equilibrium relationships). Cointegration is a relatively new statistical concept designed to deal explicitly with the analysis of the relationship between nonstationary time series. In particular, it allows individual time series to be nonstationary, but requires a linear combination of the series to be stationary. Therefore, the basic idea behind cointegration is to search for a linear combination of individually nonstationary time series that is itself stationary. Evidence to the contrary provides strong empirical support for the hypothesis that the integrated variables have no inherent tendency to move together over time.

Several methods have been proposed in the literature to estimate cointegrating vectors--see Engle and Yoo (1987) and Gonzalo (1994) for a survey and comparison. The most frequently used Engle-Granger (1987) approach is to select arbitrarily a normalization and regress one variable on the others to obtain the ordinary least squares (OLS) regression residuals e. A test of the null hypothesis of no cointegration (against the alternative of cointegration) is then based on testing for a unit root in the regression residuals e using the ADF test and simulated critical values which correctly take into account the number of variables in the cointegrating regression. This approach, however, does not distinguish between the existence of one or more cointegrating vectors and the OLS parameter estimates of the cointegrating vector depend on the arbitrary normalization implicit in the selection of the dependent variable in the regression equation. As a consequence, the Engle-Granger approach is well suited for the bivariate case which can have at most one cointegrating vector.

Table 2 reports asymptotic p-values [computed using the coefficient estimates in MacKinnon (1994)] of bivariate cointegration tests (in log levels). The entries across each row are the p-values for testing the null of no cointegration between the variable indicated in the row heading and the variable indicated in the column heading, with the variable indicated in the row heading being the dependent variable. In other words, the cointegration tests are first done with one series as the dependent variable in the cointegrating regression and then with the other series as the dependent variable-we should be very wary of a result indicating cointegration using one series as the dependent variable, but no cointegration when the other series is used as the dependent variable. This possible ambiguity is a weakness of the Engle and Granger cointegration test. The tests use a constant (but not a trend variable) and the number of augmenting lags is chosen using the AIC+2 rule mentioned earlier.

The results suggest that the null hypothesis of no cointegration cannot be rejected (at the 5 percent level), except for the pairs Rocky Mountain-San Juan, Permian-Anadarko, Anadarko-Alberta, Alberta-South Texas, and Alberta-British Columbia. That is, only five out of the twenty-eight price pairs cointegrate (at the 5 percent level and none at the 1 percent level). These results are very different from those reported in King and Cuc (1996)--there is much less cointegration across series indicating anything but an east-west split. The difference is due to the different time period than that considered in King and Cuc (1996) and possibly due to the inclusion of a trend variable in their cointegrating regressions, which reduces degrees of freedom and the power of the test--reduced power means that they conclude that the series cointegrate when in fact they don't. Notice that the King and Cuc (1996) study is not clear on how deterministic components in the time series were treated.

To investigate the robustness of these results to alternative testing methodologies, (under the assumption that North American natural gas prices are determined simultaneously) we consider the joint modelling of these prices and test for shared stochastic trends using Johansen's (1988) maximum likelihood extension of the Engle and Granger approach. Johansen's maximum likelihood approach to the estimation of the number of linearly independent cointegrating vectors for a vector autoregressive process, X., of order p involves (i) regressing [Multiple line equation(s) cannot be represented in ASCII text] on [Multiple line equation(s) cannot be represented in ASCII text] (ii) regressing [Multiple line equation(s) cannot be represented in ASCII text] on the same set of regressors and (iii) performing a canonical correlation analysis on the residuals of these two regressions--see Johansen (1988) for more details or Serletis (1994) for an application.

We search for shared stochastic price trends among prices within two price groups--eastern and western. If any shared trends are found in the eastern (western) price group [as King and Cue (1996) suggest], then they might sensibly be thought of as the eastern (western) natural gas price trends. In fact, according to King and Cuc (1996), prices within each price group tend to move together, responding to the same set of fundamentals, meaning that there is one shared stochastic price trend within each price group. Using the Engle and Granger (1987) terminology, we say that in an n-variable system with m cointegrating vectors there are n-m common trends.

Tables 3 and 4 report the results of the cointegration tests based on monthly VARs of various lag lengths for the eastern and western price groups, respectively. The results for intermediate lag length are similar. Two test statistics are used to test for the number of cointegrating vectors: the trace and maximum eigenvalue (lambda[max]) test statistics. In the trace test the null hypothesis that there are at most r cointegrating vectors where r = 0, 1, 2, and 3 is tested against a general alternative whereas in the maximum eigenvalue test the alternative is explicit. That is, the null hypothesis r = 0 is tested against the alternative r = 1, r = 1 against the alternative r = 2, etc. The 95 percent critical values of the trace and maximum eigenvalue test statistics are taken from Osterwald-Lenum (1992).

Clearly, the two test statistics give similar results in both Tables 3 and 4. In particular, the trace and lambda[max] statistics reject [at conventional significance levels, based on the critical values reported by Osterwald-Lenum (1992)] the null hypothesis of no cointegrating vectors (r = 0) and accept the alternative of one or more cointegrating vectors. However, the null of r less than or equal to 1 cannot be rejected, indicating no more than one cointegrating vector within each natural gas price group. Hence, confirming the impression from Table 2, natural gas spot prices within each price group respond to different underlying stochastic components.

When interpreting the results in terms of convergence, it should be noted that cointegration analysis cannot in principle detect convergence, because it fails to take account of the fact that convergence is a gradual and on-going process, which implies that statistical tests should lead to reject the null hypothesis of no cointegration only when convergence has already taken place-see, for example, Bernard (1992). In other words, the tests conducted here are tests for convergence over the whole period under consideration, but these tests are not tests of a move from non-convergence to convergence-the latter being the issue that King and Cuc (1996) mainly investigate.

CONCLUSION

This article explored the behavior of North American natural gas price trends and their interrelations. The degree of shared trends among natural gas prices is of considerable importance. For example, if natural gas prices share trends, in the sense that their stochastic trend components are proportional to each other, then natural gas markets have an error-correction mechanism--that is, every permanent shock in one market is ultimately transmitted to the other markets.

We applied the Engle and Granger (1987) two-step procedure to bivariate natural gas price relationships and we also tested for the number of common stochastic trends among prices within eastern and western markets using the powerful multivariate approach due to Johansen (1988). The results led to the conclusion that natural gas prices do not cointegrate and that, in particular, natural gas prices within each area (eastern and western) are driven by different stochastic trends, meaning that the east-west split does not exist.

One way to interpret these results is in terms of the absence or presence of unexploited profit opportunities. In the case of integrated price series that cointegrate, the price differential is stationary, implying price convergence, a high degree of price competition, and the absence of unexploited profit opportunities. In this case, every permanent shock in the trend of one series is ultimately transmitted to the trend of the other series. In the case, however, of integrated price series that do not cointegrate (which is the case of North American natural gas spot prices), the difference between the respective prices fluctuates stochastically, in excess of transmission and transaction costs, indicating the failure of potential arbitrage to discipline prices. In this case, the marginal value of the commodity across locations differs by more than transmission and transaction costs suggesting unexploited profit opportunities.

Table 1. Marginal Significance Levels of Unit Root Tests in North American Natural Gas Spot Prices

Market	A. Log levels		
	WS	ADF	Z(t[alpha])
Rocky Mountain	0.107	0.218	0.110
San Juan	0.030	0.127	0.065
Permian	0.010	0.044	0.020
Anadarko	0.063	0.139	0.074
South Texas	0.078	0.055	0.031
Louisiana	0.127	0.219	0.060
British Columbia	0.113	0.232	0.099
Alberta	0.092	0.274	0.158

Market	First differences of log levels WS	ADF	Z(t,)
Rocky Mountain	0.000	0.000	0.000
San Juan	0.000	0.002	0.000
Permian	0.000	0.001	0.000
Anadarko	0.000	0.000	0.000
South Texas	0.000	0.000	0.000
Louisiana	0.000	0.000	0.000
British Columbia	0.000	0.000	0.000
Alberta	0.000	0.000	0.000

Notes: Tests use a constant (but not a time trend). Numbers are tail areas of unit root tests. The number of augmenting lags is determined using the AIC+2 rule. p-values less than 0.05 reject the null hypothesis of a unit root at the 0.05 level of significance.

Table 2. Marginal Significance Levels of Bivarient Engle-Granger (1987) Cointegration Tests between North American Natural Gas Spot Prices

	Rocky Mountain	San Juan	Permian	Anadarko
Rocky Mountain	-	0.032	0.641	0.434
San Juan	0.017	-	0.930	0.122
Permian	0.185	0.627	-	0.003
Anadarko	0.142	0.048	0.018	-
South Texas	0.590	0.476	0.517	0.726
Louisiana	0.973	0.972	0.894	0.983
British Columbia	0.225	0.211	0.292	0.246
Alberta	0.226	0.259	0.443	0.048

	South Texas	Louisiana	British Columbia	Alberta
Rocky Mountain	0.720	0.978	0.215	0.008
San Juan	0.752	0.968	0.112	0.156
Permian	0.364	0.684	0.151	0.048
Anadarko	0.637	0.958	0.097	0.043
South Texas	-	0.936	0.067	0.023
Louisiana	0.971	-	0.283	0.170
British Columbia	0.286	0.572	-	0.002
Alberta	0.045	0.234	0.008	-

Notes: All tests use a constant (but not a trend variable). The number of lags is determined using the AIC+2 rule. Asymptotic p-values are computed using the coefficients in MacKinnon (1994). Low p-values imply strong evidence against the null of no cointegration.

Table 3. Johansen Tests for Cointegration among Eastern (Anadarko, Louisiana, Permian, & South Texas) Natural Gas Prices

	k = 2		k = 4		k = 6
H[0]	Trace	lambda [max]	Trace	lambda [max]	Trace
r = 0	68.068*	45.368*	46.285	32.386*	49.066
	22.699	11.007	13.898	7.528	17.850
	11.692	9.727	6.370	4.651	8.615

	1.964	1.964	1.718	1.718	2.840

Critical Values

	k = 6 lambda [max]	Trace 95%	90%	Trace 95%	90%
r = 0	31.216*	53.116	49.648	28.138	25.559
	9.235	34.910	32.003	22.002	19.766
	5.774	19.964	17.852	15.672	13.752
	2.840	9.243	7.525	9.243	7.525

Notes: Critical values are from Osterwald-Lenum (1992). k refers to the number of lags in the VAR. Drift maintained. An asterisk indicates significance at the 5 % level.

Table 4. Johansen Tests for Cointegration among Western (Alberta, British Colmbia, Rocky Mountain, and San Juan) Natural Gas Prices

	k = 2		k = 4		k = 6
H[0]	Trace	[max]	Trace	[max]	Trace
r = 0	63.323*	29.779*	64.252*	30.921*	68.283*
r < I	33.526	15.315	33.331	15.525	13.616
r < 2	18.211	13.252	17.805	13.185	15.801
r < 3	4.959	4.959	4.620	4.620	3.169

	k = 6 lambda [max]	Trace 95%	90%	Trace 95%	90%
r = 0	29.667*	53.116	49.648	28.138	25.559
	22.814	34.910	32.003	22.002	19.766
	12.632	19.964	17.852	15.672	13.752
	3.169	9.243	7.525	9.243	7.525

Notes: Critical values are from Osterwald-Lenum (1992). k refers to the number of lags in the VAR. Drift maintained. An asterisk indicates significance at the 5% level.

MAP:Figure 1. The North American Natural Gas Industry - An East-West Split?

GRAPH: Figure 2. Western Prices

GRAPH: Figure 3. Eastern Prices

REFERENCES

Bernard, Andrew B. (1992). "Empirical Implications of the Convergence Hypothesis." Working Paper, MIT, Cambridge, MA.

Beveridge, S. and C. Nelson (1981). "A New Approach to Decomposition of Economic Time Series into Permanent and Transitory Components with Particular Attention to Measurement of the Business Cycle." Journal of Monetary Economics 7: 151-174.

Dickey, David A., and Wayne A. Fuller (1981). "Likelihood Ratio Statistics for Autoregressive Time Series with a Unit Root." Econometrica 49 (July): 1057-72.

Engle, Robert F., and Clive W. Granger (1987). "Cointegration and Error Correction: Representation, Estimation and Testing." Econometrica 55 (March): 251-76.

Engle, Robert F., and B. Yoo (1987). "Forecasting and Testing in Cointegrated Systems." Journal of Econometrics 35: 143-159.

Gonzalo, Jesus (1994). "Five Alternative Methods of Estimating Long Run Equilibrium Relationships." Journal of Econometrics 60: 203-33.

Haldane, A.G. and S.G. Hall (1991). "Sterlings's Relationship with the Dollar and the Deutschemark: 1976-89" The Economic Journal 101: 436-443.

Johansen, Soren (1988). "Statistical Analysis of Cointegration Vectors." Journal of Economic Dynamics and Control 12: 231-54.

King, Martin and Milan Cuc (1996). "Price Convergence in North American Natural Gas Spot Marketa." The Energy Journal 17(2): 17-42.

MacKinnon, James G. (1994). "Approximate Asymptotic Distribution Functions for Unit-Root and Cointegration Tests." Journal of Business and Economic Statistics 12:167-176.

Nelson, Charles R. and Charles L. Plosser (1982). "Trends and Random Walks in Macroeconomic Time Series: Evidence and Implications." Journal of Monetary Economics 10: 139-162.

Osterwald-Lenum, Michael (1992). "Practitioners' Corner: A Note with Quantiles for the Asymptotic Distribution of the Maximum Likelihood Cointegration Rank Test Statistic." Oxford Bulletin of Economics and Statistics 54: 461-471.

Pantula, Sastry G., Graciela Gonzales-Farias, and Wayne A. Fuller (1994). "A Comparison of Unit-Root Test Criteria." Journal of Business and Economic Statistics 12: 449-459.

Phillips, P.C.B. (1987). "Time Series Regression with a Unit Root." Econometrica 55: 277-301.

Phillips, P.C.B., and Pierre Perron (1988). "Testing for a Unit Root in Time Series Regression." Biometrica 75: 335-346.

Serletis, Apostolos (1994). "A Cointegration Analysis of Petroleum Futures Prices." Energy Economics 16: 93-97.