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WORKING PAPER

Global LNG Pricing Terms and Revisions: An Empirical Analysis

by

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Executive Summary

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While much has been made in recent years about the increasing liquidity and size of a spot market for liquefied natural gas (LNG), with some analysts and governments even discussing creation of a centralized trading hub, most LNG is still sold under confidential, bilateral long-term contracts as has been the case since the 1960s. In fact, in 2013, according to data from the International Group of Liquefied Natural Gas Importers (GIIGNL), 73% of all LNG trades took place under long-term contracts (LTCs), which are especially prevalent in Asian markets. Despite the fact that this constitutes an enormous trade, there is very little transparency about how prices are specified, what actual transaction prices are or when pricing terms change.

Using publicly available customs data on sixteen different trade-routes of the largest importers of LNG, I apply sophisticated econometric techniques for detecting structural breaks of an unknown number and date to estimate and characterize the empirical relationship between LNG import prices and crude oil prices. These estimates allow me to make statements about the underlying pricing terms of LNG contracts, as well as when and how the pricing terms are revised. My results complement an existing literature on gas market integration, which considers cross-market convergence of prices without modeling the role of oil-indexed LTCs in determining LNG prices.

It is generally accepted that LTCs set LNG prices equal to an intercept term plus a slope times a crude oil benchmark. It is also known that some Japanese contracts have specified an "S-curve" that moderates the effect of very high or low oil prices on LNG prices. The exact parameters of the contracts, however, are not known. While it might be possible to estimate these using standard cointegration approaches, LTCs specify re-negotiation clauses, and we should expect that pricing terms may change over the course of a 20-year contract. In particular, pricing may be renegotiated when changes in market fundamentals cause the price of LNG inside a contract to diverge from the value outside the contract—the price a seller might receive from a spot sale less the transportation differential or the buyer, from an alternative supplier. Both changes in pricing terms and S-curve behavior that modifies the parameters of linear pricing terms outside of a mid-range are likely to induce a structural break in the empirical pricing relationship. Ignoring these changes not only gives biased and inconsistent estimates for contract parameters but also fails to deal with an important feature of contracts. By allowing for structural breaks of an unknown number and date, I avoid potential bias and can speak to the timing and types of revisions that occur in pricing terms.

My results strongly suggest that LNG is tightly indexed to oil, but terms are considerably more complex and varied than rules of thumb. Japanese contracts appear to have undergone the most revision. Initial breaks may be due to S-Curve behavior above thresholds at either \$25 or \$39 oil and occur in 2000 and 2004. Subsequently, tight LNG markets during the mid-2000s caused a mismatch between the price of LNG determined by the upper tail of an S-curve and its value as a substitute for oil, which steadily rose in price. Revisions brought LNG prices back to rough thermal parity with oil. It is interesting to note that pricing relationships do not undergo wholesale changes after Fukushima in 2011. This suggests that LTCs function as a form of insurance against shocks. Additionally, the fact that almost all contracts link current LNG prices to past oil prices (usually a weighted average of multiple months) may serve to smooth the magnitudes and timing of pricing shocks. LTCs in South Korea, Taiwan and Spain also appear oil-indexed but have far fewer revisions and likely no S-curve behavior. In particular, Korean and Taiwanese contracts generally set LNG prices to rough thermal parity with oil. Spanish slopes are much shallower, with LNG prices significantly below oil prices in recent years.

These results are the first rigorous characterization of global LNG pricing terms and revisions, and they complement an existing literature on international gas pricing that does not model the underlying data-generating process. This paper should be of interest to both economists and energy firms who are interested in how LNG prices are set in LTCs, under which the majority of the LNG trade is still priced. My results will also be of particular interest to firms entering into LTCs as they negotiate or re-negotiate pricing.

Global LNG Pricing Terms and Revisions: An Empirical Analysis*

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Abstract

Most LNG is sold under confidential, bilateral long-term contracts, particularly in Asia. Thus, though prices are thought to be indexed to crude oil, actual prices, contract terms and price revision clauses are not known. Therefore, I use customs data and techniques for detecting multiple unknown structural breaks in cointegrated regressions to characterize empirical pricing relationships and make inferences about pricing terms for 16 Japanese, South Korean, Taiwanese and Spanish LNG price series. LNG does appear to be indexed to oil, but terms appear considerably more complex and varied than rules of thumb. I find evidence for S-curve behavior, multiple revisions and variation in both the degree of indexation and the specification of oil benchmarks. Japanese terms are revised most. Terms for the other importers appear more stable, and indexation is weakest in Spain. This paper complements existing work on gas market integration, which largely ignores the data-generating process for LNG prices.

Keywords. LNG prices, long-term contracts, structural breaks, cointegration

1 Introduction

In 2013, 73% of all liquefied natural gas (LNG) trades took place under long-term contracts (LTCs).[1] While the spot market is substantial, most LNG is still sold under confidential, bilateral LTCs as it

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[1] Author's calculations using data from the International Group of Liquefied Natural Gas Importers (GIIGNL) report "The LNG Industry 2013." GIIGNL classifies a short-term contract as one with a term of four years or less. In 2013, spot and short-term LNG made up 27% of the total trade, which means that LTCs made up 73%.

has been since the 1960s. This means that the actual transaction prices and pricing terms are not known—possibly even between market participants—for the majority of the market. Despite the confidentiality of pricing terms, this paper shows how data analysis can be used to provide some transparency to an otherwise opaque market. Heretofore, empirical analyses of international gas prices have focused on market integration and price convergence but have not explicitly incorporated the role of oil-indexed LTCs into their modeling strategies. In contrast, I explicitly model the underlying data-generating process (DGP) and choose not to investigate statistical inter-regional relationships.

1.1 Long-term contracts and structural breaks

It is generally accepted that contracted LNG prices (particularly in Asia) are indexed to crude oil with a slope-intercept formula in which the LNG price is equal to a constant plus the crude oil price times a slope parameter (or some variation on this). One well-known rule of thumb is that LNG prices are 14.85% of crude oil prices, which is roughly thermal parity. The use of LTCs has historically made sense because it provides security for buyers and sellers in a thin market. Secure commitments can be parleyed into lower financing expenditures for the large capital investments required to liquefy or regasify LNG (Hartley, 2013). Marginal buyers of natural gas—power plants—can substitute between gas and oil products in the short term. Over longer horizons substitution is also possible on the supply side as well as in other areas of demand, such as residential heating. Thus, the contractual linkage between LNG and oil prices is simply a reflection of an underlying economic relationship (Hartley et al., 2007, 2008; Hartley and Medlock, 2014).

Pricing terms in long-term contracts are likely to be revised as market fundamentals cause the value of LNG (for example, the spot price less transportation differential) to diverge from the price specified in a LTC. Revisions to pricing formulas will alter the underlying DGP and result in a structural break in the empirical relationship. Such a situation could easily arise as supply and demand shift or new substitution possibilities arise. In fact, contracts have re-negotiation clauses that specify when and how re-negotiations occur. This is documented in Weems (2006) by a lawyer who has written a number of such sales and purchase agreements (SPAs). He quotes one contract from the 1980s with a representative renegotiation clause, which states that parties will regularly re-evaluate the contract "in good faith in the light of all the circumstances relevant at the time."

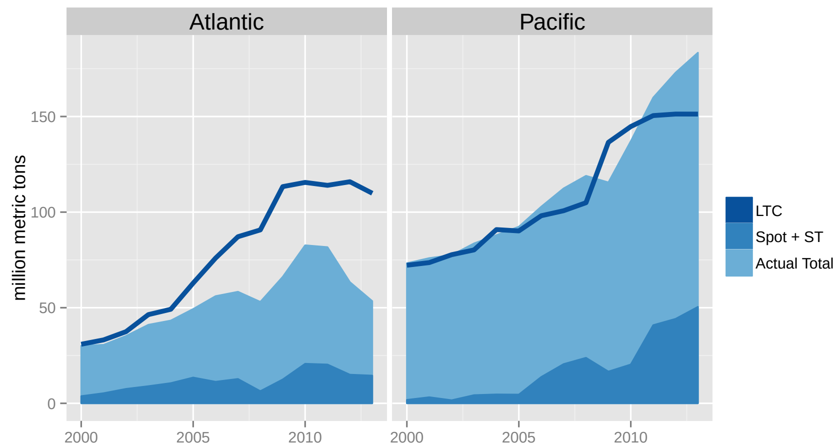


Figure 1: Actual vs short vs long-term

Later Miyamoto et al. (2009) identifies 2004 as a time when changes in market fundamentals altered the relative prices of LNG inside and outside of Japanese contracts, triggering re-negotiations:

> In 2004 .... delays to new projects and spikes in demand caused the international LNG market to begin to tighten, transforming it into an out-and-out seller's market. Consequently, this period saw moves to drastically raise LNG prices in the high crude oil price range, such as by abolishing the S-curve method and adopting a straight-line formula, and by increasing the slope of [the] pricing formula.

Over the past 14 years, LNG imports to countries in the Atlantic and Pacific basins[2] have followed different patterns that are indicative of large changes in market fundamentals particular to each region. Figure 1 plots data from GIIGNL reports from the years 2000–2013 and breaks out three quantities: 1) the realized total volume of LNG imported under any type of agreement (the light blue shaded area), 2) the volume imported under short-term or spot-contracts (the darker blue shaded area)[3] and 3) the volumes specified under long-term contracts *which are not all realized* (the darkest blue line). Since the darker blue shaded area representing spot and short-term imports is overlaid on top of the actual trade volume represented by the light blue, we can interpret the visible light blue area to be the share of LNG imported under LTCs. The darker blue shaded area is the share of spot and short-term imports.

[2] I classify the United States, Mexico and countries on the Mediterranean Sea as Atlantic importers. Gulf Coast countries are classified as Pacific importers.

[3] GIIGNL defines spot and short-term contracts as those which have a term of four years or less.

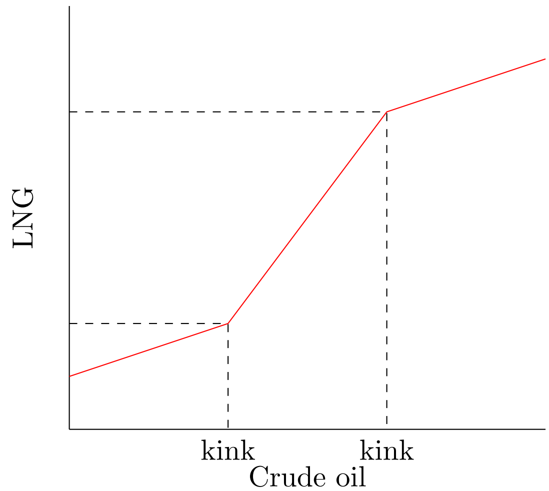


Figure 2: Hypothetical s-curve

If total imports exceed long-term commitments, we can infer that a country is purchasing extra spot volumes over and above its long-term contracts. This has clearly been the case in the Pacific since 2011, which is the date of the Fukushima nuclear disaster in Japan. The event resulted in the shut-down of nuclear plants and an increase in gas-fired electricity generation, which caused a jump in demand for LNG. If the dark blue line is above the light blue area, total imports of any sort have been less than volumes committed under LTCs. Such a situation has been true of Atlantic importers as a whole for the past decade. Since the mid-2000s, the United States has experienced large increases in natural gas production from shale, which were not expected when long-term deals would have been inked. As production of natural gas increased in North America and depressed prices, expensive LNG imports were uneconomical. One might infer that these long-term contracted volumes may have been diverted to other buyers. It is not clear *a priori* if, when or how such shifts in supply and demand will translate into LTC pricing revisions. Nevertheless, it seems plausible that they would affect contracts.

Another reason the *empirical* relationship between crude oil and LNG prices may exhibit structural changes is that a number of Asian LNG contracts specify an “S-curve” pricing formula. Such a formula dampens the impact of very high or low oil prices on LNG prices (see Figure 2 for a visual representation).[4] The presence of an S-curve implies that the slope of the relationship between oil and natural gas prices changes and becomes shallower as oil prices move into an upper or lower

[4] An S-curve could also reflect underlying economic relationships. As oil prices rise or fall, some decisions may be on a corner and substitutability between the fuels could decline.

range. Though the presence of an S-curve is not a change in the underlying DGP *per-se*, it could well appear as a break in the parameters of a linear relationship, especially since oil prices begin moving higher in 2004 and do not fall back to their previous levels. If a kink is specified at $24—as Flower (2011) claims is the case for Japan–Qatar contracts—identifying whether a break occurs in the time or price dimension would not be possible. Both Weems (2006) and Miyamoto (2008) document the use of such pricing formulas from the 1980s through mid-2000s, and filings in 1994 and 1999 related to Alaskan exports of LNG to Japan explicitly mention such S-curve behavior (see Section 3 for further discussion of the Alaskan contracts).

A final reason that breaks in the empirical LNG–oil price relationship may occur is that for a given importer–exporter pair, there can be multiple SPAs under which LNG is imported. This is more a feature of customs data I use than the long-term contracts themselves. Particularly in Japan, many companies import LNG under their own separate contracts even though the cargoes come from the same country of origin. Additionally, the proportion of spot and short-term imports—which may be priced differently than volumes purchased under LTCs—has increased substantially over the past decade. In 2000, spot and short-term volumes were just over 5% of the total trade, and in 2013, they were 27% (recall Figure 1). As the mix of SPAs under which LNG is imported changes over time, it is reasonable to expect the empirical relationship between crude benchmarks and average LNG import prices to change over time. Empirically, coefficients may shift, and model fit may worsen, which increases the estimated variance of the error process.

1.2 Empirical approach

The presence of structural breaks presents an econometric challenge. Were LNG prices determined by a single, simple slope-intercept formula, they could easily be estimated with the Johansen (1988) maximum-likelihood method or Dynamic OLS techniques proposed by Saikkonen (1991) and Stock and Watson (1993). These techniques, however, impose constant parameters, which would result in biased and inconsistent estimates in the presence of structural breaks. Furthermore, as documented by Perron (1989) and Gregory and Hansen (1996), if structural breaks are not accounted for in a cointegrating relationship, unit root tests usually suggest that the relationship is spurious when the two series are, in fact, cointegrated.

The functional form I select to represent LNG pricing formulas is a one-to-one transformation

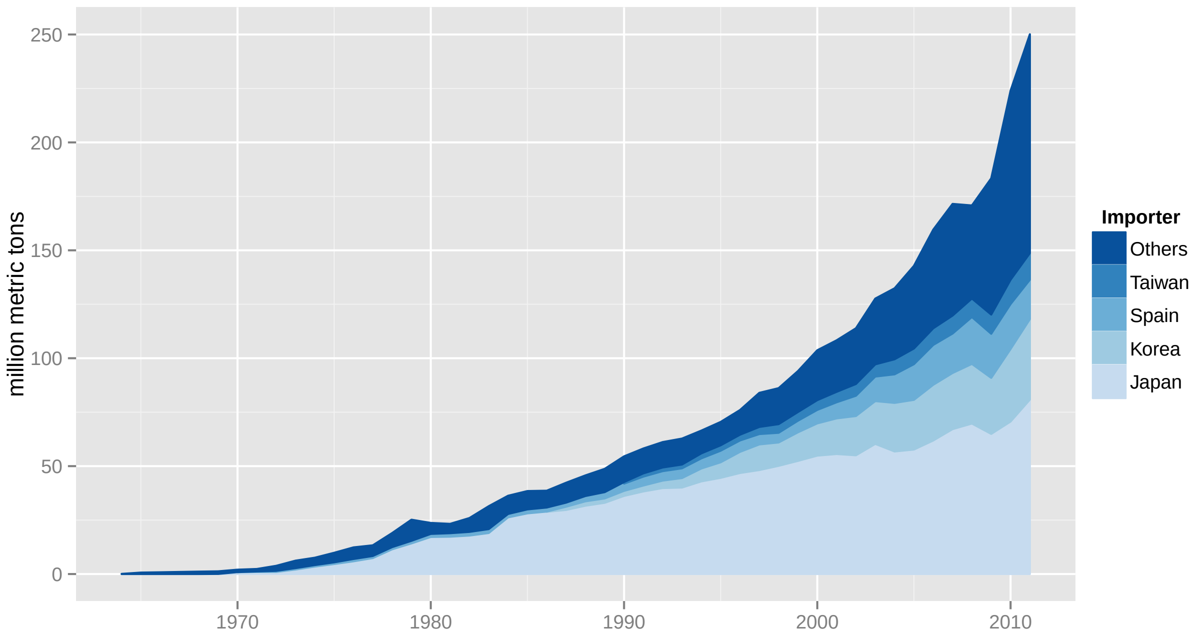


Figure 3: Historical imports. Data from Petroleum Economist.

of the pricing terms specified in long-term export license documents for Alaskan LNG. It also incorporates the rules-of-thumb for LNG pricing. Alaskan exports are explicitly priced to compete with LNG from other sources; thus, I take this form as common across all contracts. Using average prices calculated from customs data and techniques for detecting multiple, unknown structural breaks, I estimate both when and how pricing relationships change for sixteen price series from the four largest importing countries as of 2013: Japan, South Korea, Taiwan and Spain. Figure 3 plots total import volumes of these countries from 1964–2013 in a global context, and Figure 4 breaks out the import volumes associated with the price series I analyze from 2000–2013. As inspection of these plots shows, the 16 price series I analyze correspond to a large portion of global LNG imports, which means my results should be representative of broad market tendencies.

As mentioned previously, the fact that I am unable to distinguish long-term contracted and spot volumes could mean that the structural breaks I detect may not correspond to changes in pricing formulas. Nevertheless, as shown in Figure 5 (which is read in the same way as Figure 1), spot imports make up a relatively small proportion of trade volumes for many of the relationships I examine, especially in the beginning of the sample. This means that average import prices should primarily capture prices corresponding to LTCs, not spot trades.

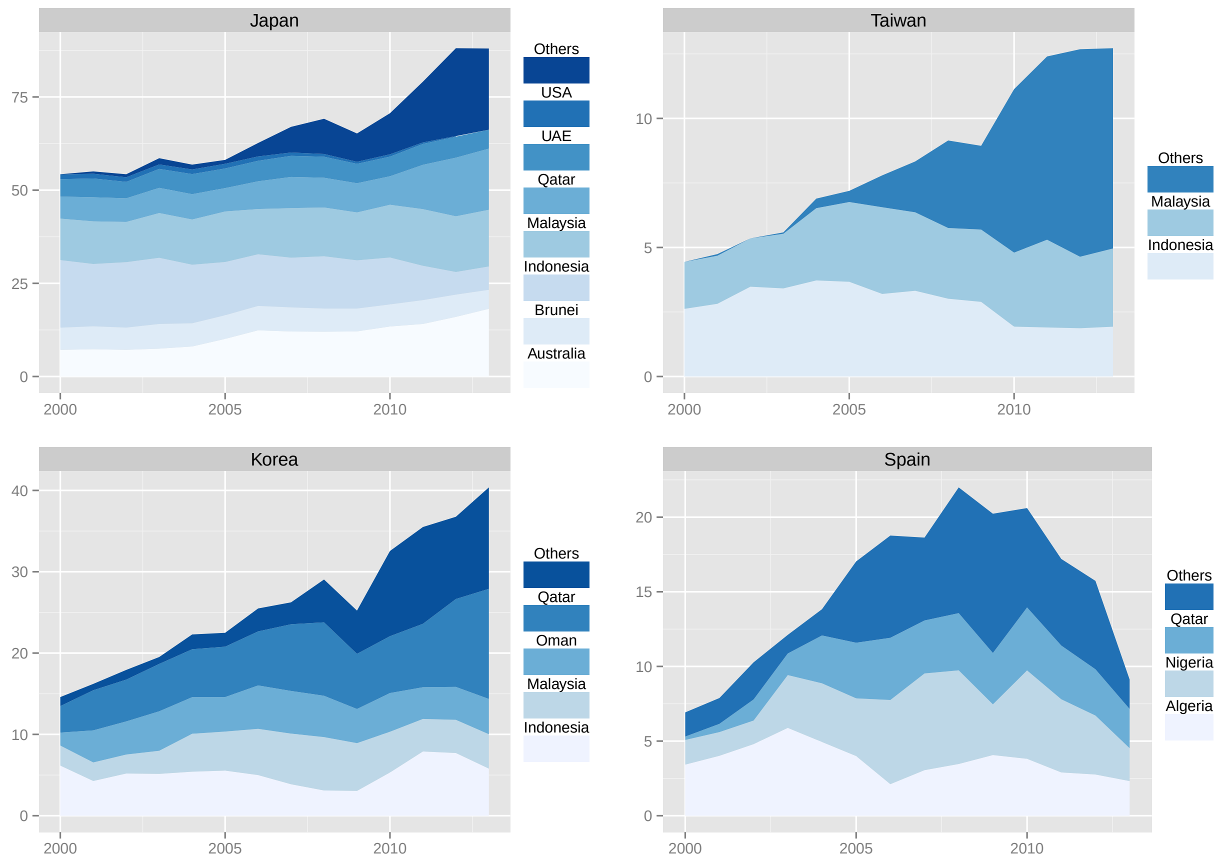
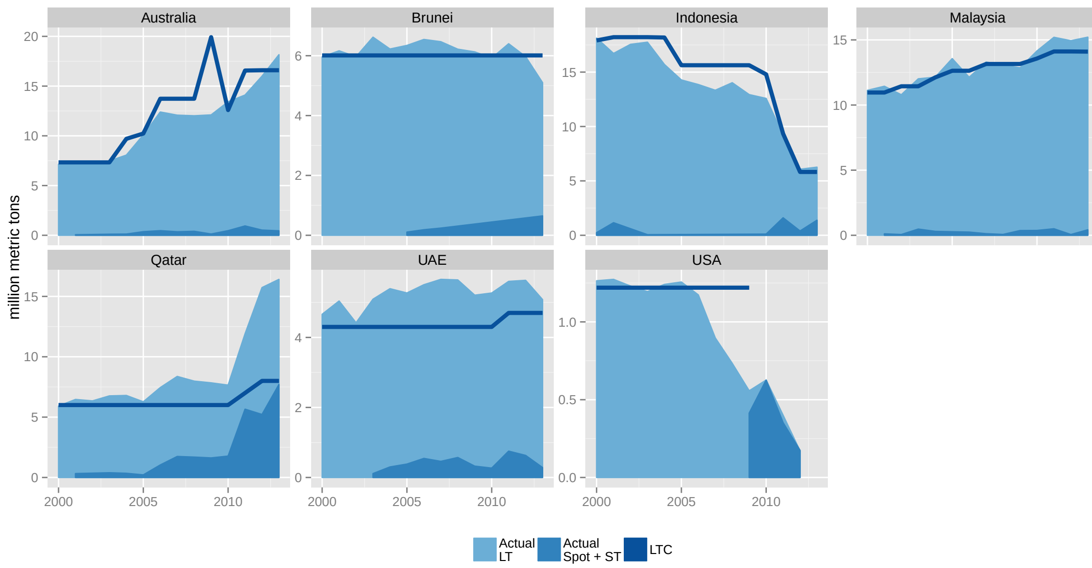
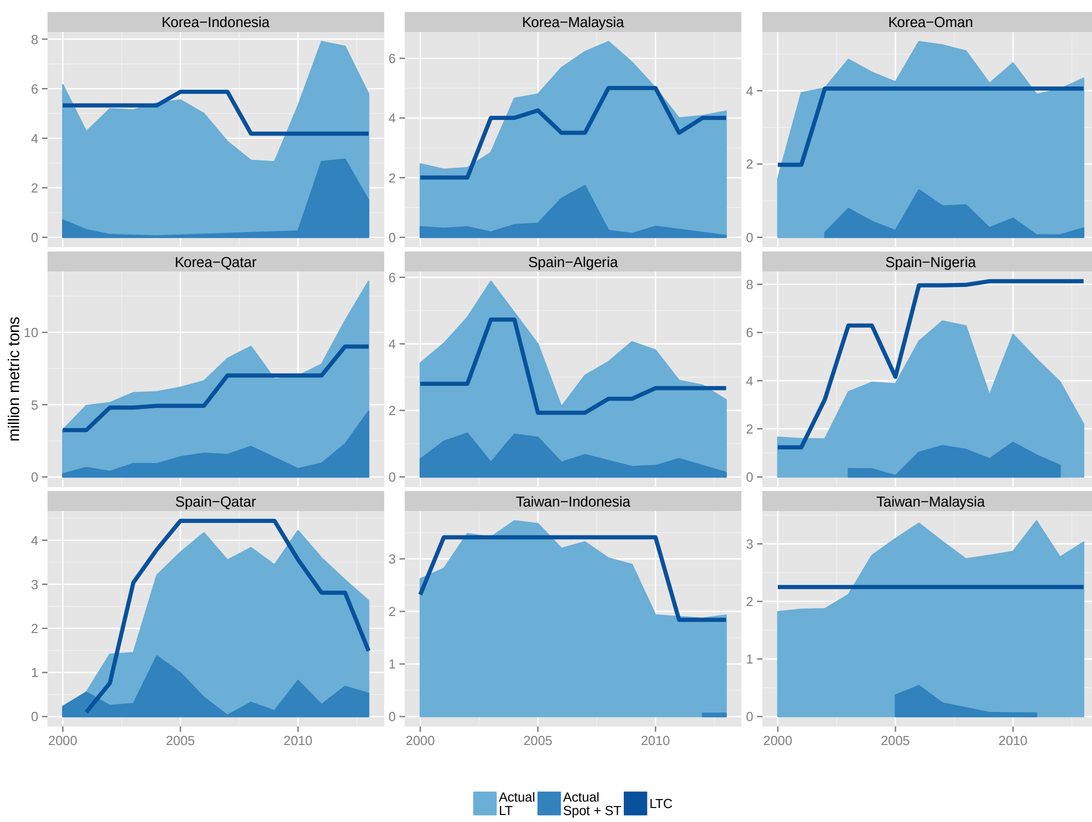


Figure 4: Import volumes associated with price series analyzed. Data from GIIGNL.



(a) Japan



(b) S. Korea, Taiwan and Spain

Figure 5: Actual, spot and contracted imports by trade route

My results suggest that Japanese contracts have undergone the most revision. Pricing terms appear to have increased as tight LNG markets during the mid-2000s caused a mismatch between the price of LNG determined by the upper tail of an S-curve and its value as a substitute for oil products. South Korean and Taiwanese LNG prices are tightly-linked to oil and seem relatively stable. Spanish contracts also appear to be oil-indexed, though less tightly and with a shallower slope.

The rest of the paper is organized as follows. Section two reviews relevant literature on global natural gas markets. Section three provides a brief overview of the econometric model employed. Sections four and five describe the data used and estimation method, respectively. Section six contains results, and section seven concludes.

2 Literature review

One strand of literature on gas prices addresses the convergence of regional spot markets to a single market, which is fundamentally different than my interest: the pricing terms in *long-term contracts* and their *revisions*, which, by definition, requires modeling the pricing terms and discrete changes in them. De Vany and Walls (1993) pioneers this literature and uses cointegration analysis to determine the degree of spot-market integration of natural gas prices. In particular, the authors find that unbundling pipeline ownership and capacity rights in the United States increased market integration. More recently, Siliverstovs et al. (2005) uses the Johanssen procedure to investigate relationships between European, Japanese and North American gas prices. The authors find that European and Japanese gas markets are cointegrated but reject an absolute version of the Law of One Price, which would require the cointegrating vector to be $\begin{pmatrix} 1 & -1 \end{pmatrix}$. The authors also reject the hypothesis that North American prices are cointegrated with prices in Europe or Asia. The authors do not consider structural change or distinguish between trading partners as I do. Neumann et al. (2006) and a subsequent paper, Neumann and Cullmann (2012) use a time-varying Kalman filter to understand both the degree to which daily European spot prices have converged and the process of convergence. Their approach is more flexible than in previous studies and allows the degree of convergence to evolve over time. They find that the degree of market integration across Europe has increased.

Most recently, Li et al. (2014) use tests for price convergence among multiple series to examine the convergence of gas prices during the period 1997–2011 in South Korea, Taiwan, and Japan as well as the Henry Hub and NBP benchmarks. The authors' objective is to see how (and whether) markets integrate and prices converge such that the Law of One Price is satisfied. (As stated, the functional form used ignores transportation differentials.) Thus, they smooth all Asian prices by aggregating over exporters when testing for convergence, taking logs and filtering out high-frequency movements with the Hodrick-Prescott filter. The integration process they measure is, by construction, smooth. In contrast, my interest is in characterizing the underlying DGP for each price series over a longer time period. Thus, I work with disaggregated, untransformed and unfiltered series. I estimate the parameters of underlying pricing formulas (including the lag structure) as well as the distinct dates when the pricing regimes change. Li et al. (2014) finds some evidence that Japanese prices join a South Korea–Taiwan "convergence club" only in later years, and they suggest that the later entry of Japan is due to lags in LTC revisions. My results provide an explanation for these authors' finding. The use of S-Curve pricing in several Japanese LTCs meant that during the mid-to-late 2000s, Japanese LNG prices were determined by the upper tail of S-Curve, which meant Japanese prices were significantly discounted in relation to South Korea and Taiwan, as well as crude oil on an energy-equivalent basis. In 2008–2009, I find that the S-Curves are eliminated, and prices regain rough thermal parity. My paper complements the market-integration literature since the latter focuses on cross-country pricing relationships while ignoring the underlying terms that set prices. I take the underlying data-generating process seriously but ignore the statistical cross-country pricing relationships.

A number of theoretical works—Brito and Hartley (2007), Ikonnikova et al. (2009) and Hartley (2013)—have examined the interplay of the spot market and bilateral contracts, ignoring the role of contract revision. Brito and Hartley (2007) considers contracts in a search model and captures the idea that the attractiveness of long-term contracts depends on expectations about how thin the market will be for short-term trades. The paper suggests that the formation of a spot market could be quite rapid as liquidity increases and parties can expect to find trading partners easily. Ikonnikova et al. (2009) approaches the question of long-term contracts from a game-theoretic approach. The authors incorporate the emergence of destination-flexibility in their model and find that more liquid spot markets and greater uncertainty about demand lead to more flexible contracts and, hence, more spot trading. Hartley (2013) examines the question of why buyers and sellers might want to have a long-term contract at all. Hartley finds that long-term contracts are beneficial primarily insomuch as they lead to reduced project financing costs (in the parlance of industry, these are “bankable contracts”). Contracts are less beneficial when the average difference between the net-back prices faced by exporters and the spot-market prices faced by their buyers—the rent which both parties split in a LTC—decreases. As the spot market deepens, firms should also increase their use of flexibility in contracts and decrease reliance on long-term contracts. The importance of stabilizing cash flows forms the basis of predictions in Mitrova (2014) that Russia will continue to sell natural gas almost exclusively in long-term contracts with strong oil-indexation. Mitrova reasons that new gas supplies will be very expensive to develop, so securing buyers is necessary to finance the projects.

3 Model and breakpoint estimation

The United States Federal Energy Regulatory Commission (FERC) provides copies of three amendments and FERC approvals (in 1988, 1991 and 2000)[5] to the long-term LNG sales and purchase agreement under which Phillips and Marathon exported LNG from Kenai, Alaska to the Tokyo Electric Power Company (TEPCO) and the Tokyo Gas Company in Japan. These documents list three different pricing regimes. Each regime specifies a “redetermination of the contract price” when the base crude index (Japan Crude Cocktail, in this case) is outside of the range \$13–26 (1994) and subsequently \$11–25 (1999). Such “redetermination” could certainly be consistent with S-Curve behavior, and the two thresholds are close to those specified in Qatar–Japan contracts as described by Flower (2011). Alaskan LNG-related documents describing the pricing terms all explicitly mention that LNG is priced to be competitive with both crude oil as well as LNG imported from other countries. This is evidence that the functional form of the Japan–USA contract should be common across other contracts as well.

Slope-intercept LNG pricing formulas that change over time can be expressed as

$$y_t = c_j + \delta_j \tilde{z}_t + u_t \qquad t = T_{j-1} + 1, \ldots, T_j \tag{1}$$

[5]Available at `http://www.fossil.energy.gov/programs/gasregulation/authorizations/Authorizations.html`

where $y_t = P_{LNG,t}$, $z_t = P_{Oil,t}$ and $\tilde{z}_t = \sum_{r=0}^{p+1} \theta_r z_{t-r}$ is a weighted average of present and past oil prices. The variables y_t and z_t have unit roots, and u_t is a stationary, mean-zero process. Time is indexed by t with breakpoints at T_j and break fractions $\lambda_j = T_j/T$. This means that there are $j = 1, \ldots, m$ breaks and $m+1$ regimes. I follow the convention that $T_0 = 0$ and $T_{m+1} = T$. The breakpoints T_j are unknown, and I wish to estimate them along with c_j, θ_{rj} and δ_j. Using three facts

$$\tilde{z}_t = \theta_0 z_t + \sum_{r=1}^{p+1} \theta_r z_{t-r} \qquad z_{t-r} = z_t - \sum_{s=0}^{r-1} \Delta z_{t-s} \qquad \sum_{r=0}^{p+1} \theta_r = 1,$$

equation (1), which bases LNG prices on a weighted average of oil prices in $t, t-1, \ldots, t-(p+1)$, can be easily transformed into a unique Dynamic OLS (DOLS) model (regime subscripts j are omitted for simplicity):

$$y_t = c + \delta z_t + \sum_{s=0}^{p} \pi_s \Delta z_{t-s} + u_t \tag{2}$$

where $\pi_s = -\delta \sum_{r=s+1}^{p+1} \theta_r$. The original weights can be recovered[6] as

$$\theta_0 = 1 + \pi_0/\delta \qquad \theta_{p+1} = -\pi_p/\delta \qquad \theta_{r=1,\ldots,p} = (\pi_r - \pi_{r-1})/\delta.$$

I prefer to use a DOLS model for three reasons. First, the most interesting parameter is δ, the slope in the cointegrating relationship. By including only one level of oil prices (which have a unit root), I avoid high collinearity between lagged levels and can obtain more precision in my estimates of δ. Second, asymptotic theory has been developed for the DOLS model, and the distribution of breaks is dependent only on the number of I(1) regressors, which means breaks should converge faster than they would with an Autoregressive Distributed Lag model that uses lagged levels of z_t. Finally, the transformation explicitly incorporates the restriction that the weights of past lags of oil prices sum to one.

I estimate and test both the positions and number of breaks using theory developed by Bai and Perron (1998, 2003) for a stationary context and extended by Kejriwal and Perron (2008b, 2010) to a cointegrated context. The basis for breakpoint-testing is the sup Wald test for k unknown breaks

[6] This follows from the following three facts:

$$\pi_p = -\delta\theta_{p+1} \qquad \pi_q - \pi_{q-1} = \delta\theta_q \text{ for } r = 1, \ldots, p \qquad \theta_0 = 1 - \sum_{r=1}^{p+1} \theta_r.$$

in the set of coefficients. I denote the test statistic $F_T(k)$. Given a minimum allowable segment size ϵ (which is in terms of percentage of the sample), the procedure finds the vector of break-dates that, when allowed, minimizes the sum of squared residuals (SSR) and tests the significance of this optimal break. Minimization of the SSR is equivalent to maximizing $F_T(k)$. Initial detection of *any* breaks is done with the double-maximum $UD\max F_T(M)$ test of the null hypothesis of no breaks against the alternative of an unknown number of up to M breaks. The test-statistic is simply $\max_{1\leq m\leq M} F_T(m)$. If I reject the $UD\max F_T(M)$ test, the number of breaks is chosen using sequential tests $F_T(k+1|k)$. These test the null of k breaks against the inclusion of an additional break during one of the $k+1$ segments defined by the k breaks. The point at which the sequential test no longer rejects the null hypothesis indicates the true number of breaks. Descriptions of these tests and the robust (to serial correlation) versions I use are in Appendix A. I denote the robust versions of the tests using an asterisk: $F_T^*(k)$, $UD\max F_T^*(M)$ and $F_T^*(k+1|k)$.[7]

4 Data

LNG prices are all constructed using country-level customs data on the quantity and value of imports. Prices are computed by dividing the total weight of LNG imports by the total value of the imports, and I convert the per-ton price to a per-mmbtu basis.[8] Currency conversions are required for Japan, Taiwan and Spain. Monthly averages of weekday exchange-rates from the Federal Reserve "Foreign Exchange Rates" tables were used to convert currencies to US dollars.

Japan's dataset is the longest, stretching back to 1988. While data are available at the port-level, I elect not to use a panel approach and simply aggregate to the national level. Many ports have missing months, and scatter-plots of the data are qualitatively very similar to those in Figure 7. It is unclear what treating each port separately and estimating a large number of regressions would add. Also, econometric theory has not been developed to deal with a panel of cointegrated series that has multiple, unknown structural breaks for each cross-sectional unit, so I do not take that approach, either. South Korean customs data begin in 1995. Taiwanese customs data begin in 2000. Spanish

[7] I did calculate $UD\max F_T(M)$, $F_T(k)$ and $F_T(k+1|k)$, but values were implausibly large, and residuals appear exhibit substantial persistence, which violates the assumptions of the tests.

[8] An approximate conversion is 51.813 mmbtu/metric ton. There are approximately 5.8 mmbtu per barrel of crude oil. While GIIGNL reports that list average LNG characteristics show that different liquefaction plants have different calorific contents, this data is unavailable before 2003 and may not correspond to what is traded under LTCs. In fact, the tables are for "non-contractual" characteristics.

customs data are also available at the port-level back to 1995. For the reasons above, I choose not to analyze port-level data for Spain, either. Japanese and South Korean pricing relationships were estimated with data through June 2014, while Spanish and Taiwanese relationships were estimated with data through May 2014. Summary statistics are in Table 1. The table includes Dickey-Fuller tests for the null hypothesis of a unit root. A unit root cannot be rejected at the 10% level for any series, so cointegration is an appropriate framework.

There are two logical choices for a crude oil benchmark: Brent crude and Japanese Crude Cocktail (JCC).[9] The first is available beginning in 1984, and the second, January 1988. Brent is a light, sweet and primarily waterborne crude oil from the North Sea, and it is the most commonly referenced global crude oil benchmark. JCC is a weighted average price of crude oil imports to Japan and is generally thought to be the benchmark used in Asian pricing formulas. Monthly Brent crude prices were gathered from the United States Energy Information Agency (EIA). I construct my own Japanese Crude Cocktail (JCC) from customs data since the official one is only available beginning in 2000.[10] Some LNG trade-routes contain months with no trade and, therefore, no prices. In these cases, I use the largest contiguous series of prices. Imputing missing data is always risky, and missing data may indicate a change in import-type.[11]

5 Estimation

LNG pricing formulas for Japan–USA use a weighted average of past oil prices. For other series, cross-correlograms of differenced data bear this out, showing that LNG prices in period t are usually more correlated with lagged oil prices than contemporaneous prices. Therefore, rather than pick the crude oil benchmark in time t that is most highly correlated with an LNG price series in time t, I regress LNG prices against lags zero through six of crude oil prices and choose the benchmark that yields the best fit in terms of the SSR. As expected, the Japanese, South Korean and Taiwanese

[9] According to Miyamoto et al. (2009) the Japan–Indonesia contract uses the government-reported Indonesian Crude Price (ICP), which is not available publicly.

[10] My series is a weighted average price of all crude oil imports into Japan (all HS codes beginning in 2709). I convert from Yen/kL to USD/bbl using official volume-to-weight conversions from the Petroleum Association of Japan and foreign exchange rates from the Federal Reserve. The series is quite close to the official figures, and differences appear to be primarily due to different foreign exchange rates.

[11] For example, early Korean–Australian imports, after which there are missing months, appear to be priced much higher versus Brent Crude than subsequent imports, possibly because they are all spot sales according to data from GIIGNL.

LNG prices are all most highly related to the Asian JCC benchmark (with the notable exception of Indonesian imports in South Korea and Taiwan). Spanish imports of LNG from Algeria appear to be pegged to Brent crude, but Qatari and Nigerian imports appear to be linked to JCC. Since Spain competes with Asian buyers for Qatari cargoes, it is not surprising that the benchmark would be common with Asian buyers. Brent crude is from the UK, and Algeria sells large quantities of gas to Europe. Thus, the Brent–Algerian LNG relationship is not surprising. It is unclear why Nigerian LNG—an Atlantic cargo—would be pegged to JCC. The SSRs from the two regressions and most correlated lag are listed with summary statistics in Table 1.

I determine the appropriate number of lagged differences of oil prices (the parameter p) to include in equation (2) by using the Bayesian Information Criterion (BIC), as suggested by Kejriwal and Perron (2008a). I take rejection of the $UD \max F_T^*(M)$ statistic at the 1% level to indicate the presence of at least one break. The number of breaks corresponds to the value k at which the sequential test, $F_T^*(k+1|k)$, is no longer rejected at the 1% level.[12] In most cases, the number of breakpoints selected corresponds to the minimum of the LWZ information criterion suggested by Bai and Perron (1998, 2003) and Kejriwal (2008).[13] (See Appendix A for details on the calculation.) Breakpoint test statistics are available in Tables 2, 3 and 4. In the bottom of each table, the rows BIC, LWZ, SSR_k, F_T^*, $F_T^*(k+1|k)$ and $\min(\Delta T_i)$ correspond to the Bayesian Information Criterion and LWZ information criterion (see Appendix A for details), the sum of squared residuals when k optimal breaks are allowed, the sup-Wald test for 0 breaks verus an alternative of k breaks, the sequential test of $k+1$ breaks versus the null of k breaks and the minimum length among all $k+1$ regimes corresponding to the k breaks. Equation (2) and associated test statistics are then estimated separately for each regime. I estimate the long-run variance of u_t using an Andrews (1991) HAC estimator and use this to calculate standard errors.[14] Finally, I calculate the implied

[12]Unfortunately, the minimum break-sizes I admit in the Japan–Indonesia and Japan–USA regressions are smaller than those for which the authors provide critical values. A number of Indonesian contracts expire in 2011, and the US export license changes from long-term to short-term in 2009. Allowing these breaks requires $\epsilon < 0.15$. In this case, I use the LWZ criterion, the minimum of which corresponds to a sharp decrease in the $F_T^*(k+1|k)$ test that would indicate a break were $\epsilon \geq 0.15$.

[13]Breakpoint estimation is done with a modified **R** package **strucchange**. Critical values are taken from the online appendix of Kejriwal and Perron (2010) and correspond to case a. They are available at `http://people.bu.edu/perron/papers/cv_eps152025.pdf`. Unfortunately, critical values for the case where $\epsilon = 0.10$ are not provided, so I cannot use the $F_T^*(k+1|k)$ test for Japan–Indonesia and Japan–USA prices. In both cases I used the LWZ criterion, which selected the breakpoints as would have been selected were the $\epsilon = 0.15$ critical values to have been used.

[14]The variance matrix is $\hat{\sigma}^2_{LRV,i}(X_i'X_i)^{-1}$ where $\hat{\sigma}^2_{LRV,i}$ is the estimator of the long-run variance as suggested by Andrews (1991). This is the same as equation (4) except that here $u_t = \tilde{u}_t$ (i.e., we use the residuals from segment i of the model with breaks). See Appendix A for more details.

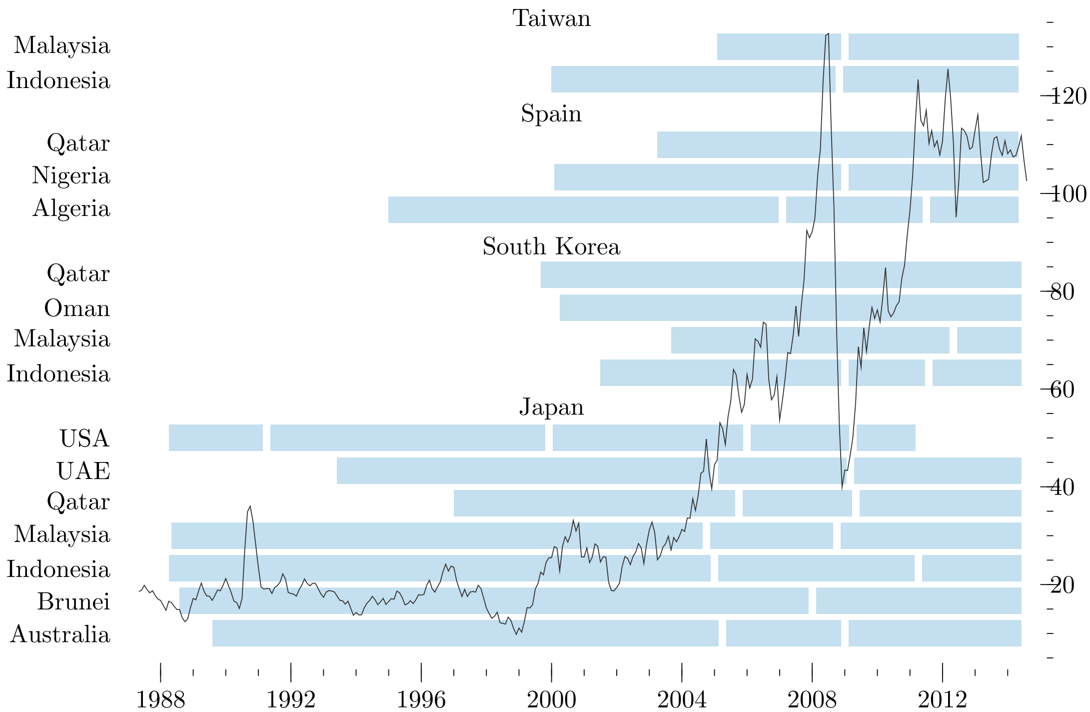


Figure 6: Contracts with breaks (plus Brent crude)

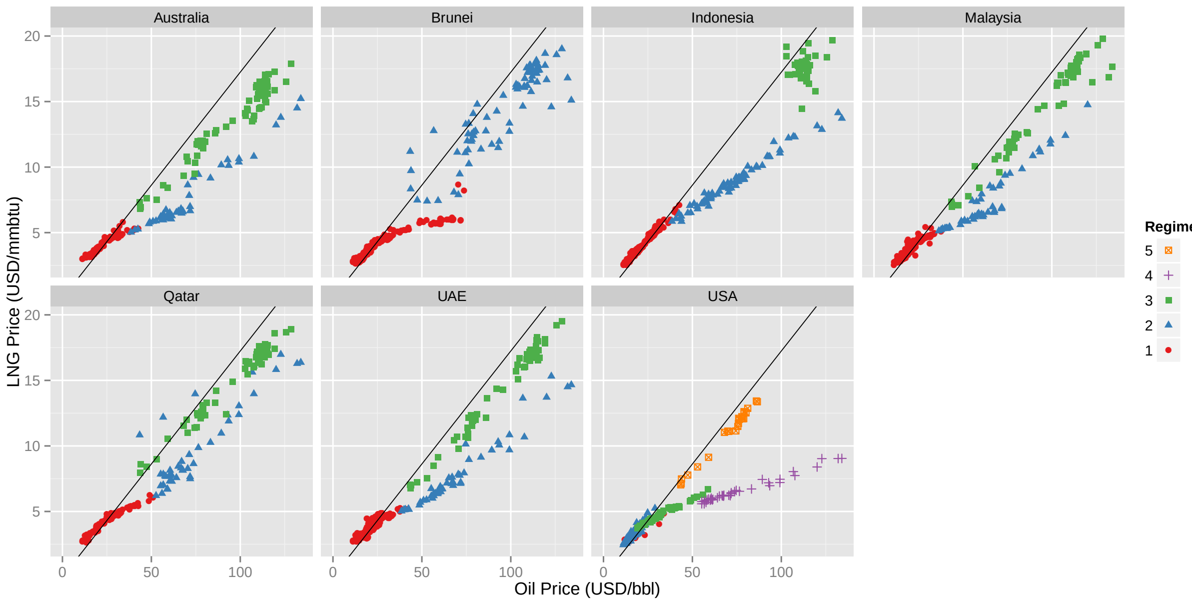
parameters for an LNG pricing formula—equation (1)—and compute standard errors using the the delta-method. These estimates of average pricing formulas are given in Tables 5, 6, 7 and 8.

6 Results

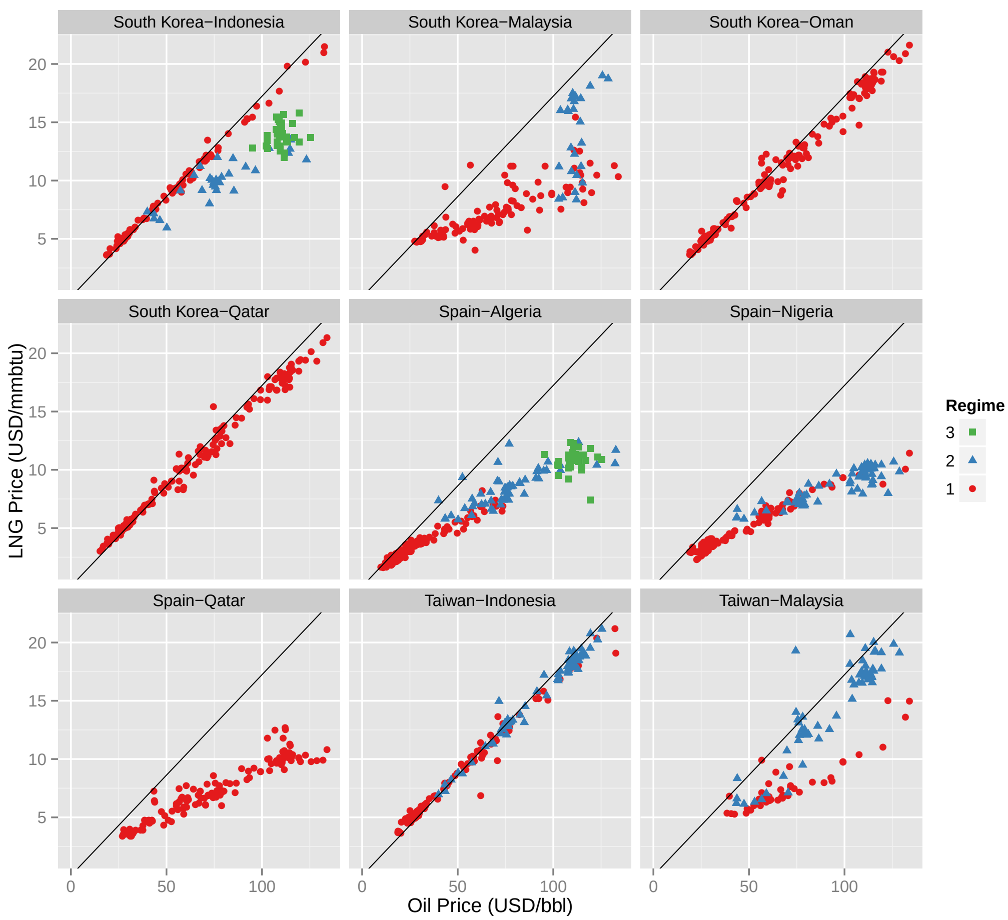
Figure 6 represents each price series as a light blue bar, and estimated breakpoints appear as white, vertical breaks. Note that the length of the bar corresponds to the length of the series I use, not the full history of trade. Brent crude prices are overlaid on top of the diagram for reference. Figures 7a and 7b depict the estimated pricing formulas in the price dimension. The vertical axis is LNG prices, and each regime is coded by color and symbol shape. The horizontal axis corresponds to the lag of the crude benchmark most highly correlated with each LNG price series. (Table 1 lists the benchmark and lag plotted.) The solid diagonal line represents rough thermal parity.[15] The same information is plotted in the time-dimension in Figures 8a and 8b.

Before proceeding to examine individual price series, three general comments are worth making. First, there is considerable variation in the lags used to form oil price benchmarks $\tilde{z}_t$. Averaging

[15]Thermal parity is approximately 1 mmbtu $\approx$ 0.172 bbl crude since 1 bbl $\approx$ 5.8 mmbtu.

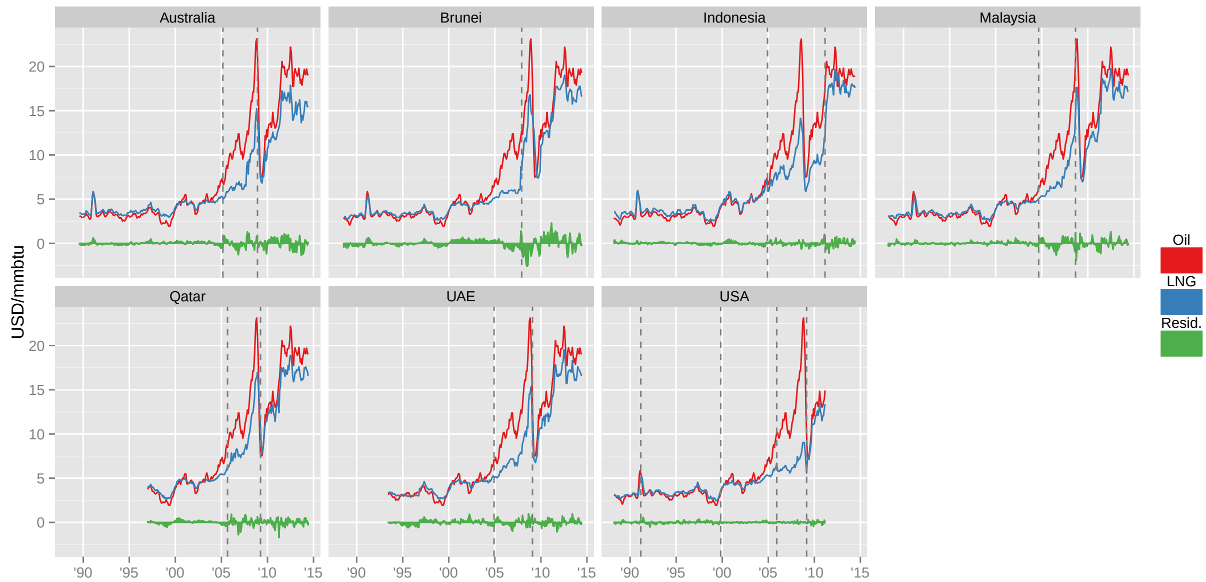


(a) Japan

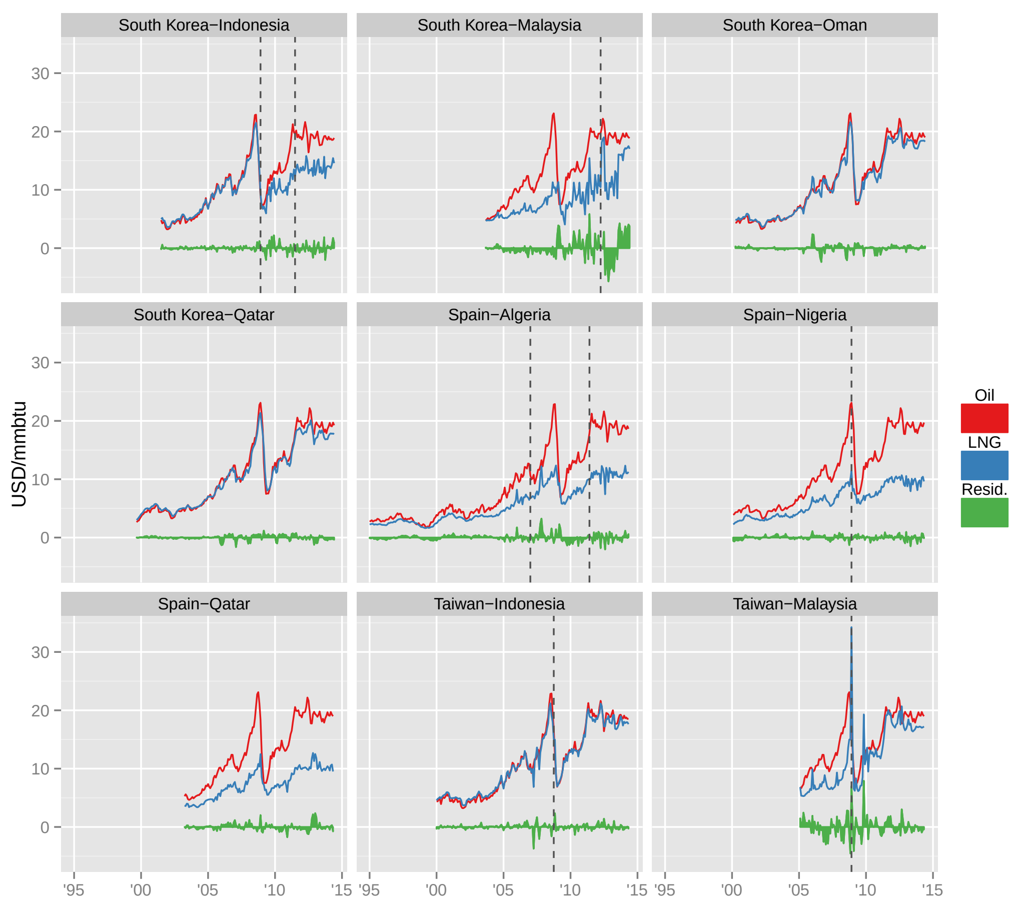


(b) S. Korea, Taiwan and Spain

Figure 7: Scatter plots of LNG versus lag of oil with maximal correlation



(a) Japan



(b) S. Korea, Taiwan and Spain

Figure 8: Time-series plots of LNG versus lag of oil with maximal correlation

over several lags should serve to dampen volatility of contracted LNG prices, but it is unclear why some pricing formulas use longer lags of oil prices than others. In general, estimates of average pricing formulas have one or two particularly significant θ_r parameters which should correspond to the lags primarily used in pricing LNG. Significant pricing lags range from the contemporaneous (θ_0) to the sixth (θ_6). Most parameter estimates for these significant lags lie in the interval $[0, 1]$ as expected, though some regressions with poorer fits have significant parameters well outside of this range.

Second, the link with oil prices is very strong, and Dickey-Fuller tests for stationarity of the residuals from equation (2) all exceed 1% critical values for the Engle-Granger cointegration test when the estimated breaks are allowed. This means that despite increases in the share of spot volumes, average LNG prices are still cointegrated with crude oil prices.

Third, it bears reiterating that breakpoints are estimated under the assumption of constant variance under the null hypothesis of no breaks using an asymptotic framework that relies on the interval between observations going to zero. Thus, breakpoint estimates should be interpreted with some caution. For instance, casual inspection of the Japan–Brunei relationship in Figure 7a suggests that the slope in the initial regime (red circles) changes as oil price rises and that the variance may be higher in the second period.[16] Thus, the true number of breaks may be over or under-estimated. Because of the changes in variance over time, I choose not to calculate confidence intervals for breakpoints even though Kejriwal and Perron (2008b) explains how to do so.

[16] Breakpoint tests do not show that Brunei has a statistically significant break that corresponds to an S-curve (which would occur within the initial red-dot regime) when the entire series is used to form test statistics. However, the variance in the first segment is much smaller than the variance in the second, so breakpoint tests might have missed a kink corresponding to an S-curve. (Such a possibility vanishes, of course, as the interval between observations goes to zero.) If only the first segment from 1988(8)–2007(12) is tested for a break of a minimum size $\epsilon = 0.15$, the $UD\max F_T^*(M)$ and $F_T^*(k+1|k)$ tests reject the null of no breaks in favor of a single break at 2003(12) at the 1% level. The break corresponds to a shallow upper tail of an S-curve which could begin in the \$30.55–\$31.07 range. (The endpoints correspond to average JCC prices in 2003(12) and 2004(1), respectively). Rather than a single slope parameter of 0.064, the two slope parameters would be 0.119 and 0.048. While the resulting plots in Figure 9, which appears in Appendix B, are visually quite convincing, this is an *ad-hoc* test, so I do not report the results in tables. Nevertheless, it is illustrative. In particular, such a procedure violates the asymptotic framework upon which the tests are based, specifically the notion that the interval between observations shrinks and the duration of the series is fixed. Graphs suggest that Malaysian and Qatari prices might also exhibit undetected S-curve behavior.

6.1 Japan

Japanese pricing relationships appear to have undergone more changes than those of other importers.[17] Figure 6 shows two clusters of breaks in Japanese contracts during 2004/2005 and 2008/2009. In particular, initial pricing regimes as plotted in Figure 7a seem to hew relatively close to thermal parity since the red circles fall along the diagonal line. When oil prices head above $25, however, LNG prices do not rise as fast and therefore fall below thermal parity. This behavior, which is suggestive of S-Curve kinks in the LNG–oil pricing relationship, corresponds to time-breaks in some series (particularly imports from Australia, Indonesia, Malaysia and UAE) in 2004 and 2005. In later regimes, LNG prices move higher and appear to regain rough thermal parity with oil prices, which roughly corresponds to the 2008/2009 cluster of breaks (evident in Figure 6). Such a narrative could correspond to the elimination of S-curves or an upward revision in pricing terms, as suggested by Miyamoto et al. (2009), though perhaps in later years than the author suggests. This is likely the round of breaks that Li et al. (2014) states brought Japan into the Asian "convergence club."[18] It is interesting to note that many series do not have breaks after the Fukushima incident since, as Medlock (2014) notes, spot prices for Asian LNG jumped dramatically immediately following the tsunami. A notable exception is the Japan–Indonesia trade, which displays a clear break precisely in March 2011. The date, however, is coincident with the expiry of a large portion of the Japan–Indonesia contracts and a sharp fall in volumes, not just the Fukushima disaster. Figure 5a shows a large drop in both actual total volumes and contracted volumes from Indonesia. The stability of pricing formulas after the Fukushima incident suggests that LTCs function as a form of insurance against unexpected shocks. This is because LTCs tie prices in a relatively thin market (LNG) to a deep, liquid market (crude oil). A quantity shock to one country in the very deep crude oil market may not represent a relatively large change in global crude oil volumes; however, a shock to one country in the LNG market is likely to represent a much larger disruption proportionally and induce more volatility in prices. Thus, LNG contracts do not

[17] Japan is the largest importer of LNG with the largest number of contracts, buyers and regasification terminals. Thus, the larger number of breaks could be due to a changing mix of SPAs underlying the weighted average import price. However, GIIGNL conract data suggests that Japanese buyers often negotiate contracts together, which should alleviate issues of underlying heterogeneity in contracts.

[18] The 2008/2009 period was also a time of great upheaval in financial markets and saw extraordinary swings in oil prices. I use 1% critical values in part to avoid the possibility that breaks are due to this volatility. Additionally, there are very distinct changes in the estimated slope parameters, $\hat{\delta}$, which is evident from Figure 7a

function as a device for a monopolistic seller to opportunistically extract rents but as a mechanism for counterparties to stabilize cash-flows and split rents that arise from relative transportation differences.

One explanation for the larger number of Japanese breaks is that only Japanese pricing formulas appear to exhibit S-curve behavior. Since oil prices rose steadily in the mid-2000s, an S-curve could be identified as a break in time, not only the range of oil prices. Two series exhibit very clear S-curve behavior that corresponds to breaks in time: USA and Indonesia in 1999(11) and 2004(12), respectively. In each case, the estimated slope parameter, $\hat{\delta}$, falls by almost 50%. When the USA break occurs between 1999(11) and 1999(12), average JCC prices were \$23.25 and \$24.94. They do not fall below \$25 again except for a brief period 2001(11)–2002(04). The Kenai export license specifies a kink precisely in the \$25–\$26 range.[19] The Indonesia break at 2004(12)–2005(1) corresponds to average JCC prices of \$39.82 and then \$38.54, which would mean that the Indonesia kink is higher.

It is possible that some breaks are the result of the signing of new LTCs or the expiry of existing ones, not revisions to pricing formulas. Certainly this may be the case for Indonesia. By the same token, however, according to GIIGNL data, Japanese imports from Brunei are governed by a single LTC from 1993–2013. Figure 5a shows that import volumes are relatively stable and include limited spot volumes beginning only in 2005. Thus, it would appear that breaks in the Japan–Brunei pricing relationship are driven by changes to the pricing terms specified in the contract, not heterogeneity in underlying contracts. I take the case of Brunei as evidence that breaks are not all the result of underlying heterogeneity in LTCs.

6.2 South Korea, Taiwan and Spain

South Korea, Taiwan and Spain are all major importers of LNG, though they purchase less than Japan. Only six of nine series exhibit breaks, and there are fewer breaks per relationship. (Just two series have more than one break.) The pricing relationships and, therefore, underlying pricing formulas, appear to be more stable. Furthermore, slope estimates—$\hat{\delta}$ in Tables 7 and 8—and scatter-plots in Figure 7b suggest that revisions have tended to consist of relatively minor modifications

[19] Similarly, Flower (2011) states that SPAs between Qatar and Japan use an S-Curve with an upper kink at \$24/bbl. For these contracts the slope parameter is 0.1485 in the mid-range, and in the tails it is 0.07.

to slopes and lags used in constructing moving average prices. Given that Taiwan and Korea have a single buyer (KOGAS and CPC, respectively) and that a number of exporting countries have a single company selling LNG, not multiple exporting firms, it seems reasonable to expect there to be less heterogeneity in the SPAs underlying average import prices.

Slopes for South Korean and Taiwanese pricing relationships are relatively steep (with the exception of imports from Malaysia and later South Korea–Indonesia imports). Slope parameters for these contracts appear to be similar, and estimates stay in a relatively narrow range between 0.152 and 0.158, which is substantially higher than Japanese values and closer to both thermal parity (0.172) and the 0.1485 rule-of-thumb. The second South Korea–Indonesia break in 2011(7) may be the result of large increases in spot volumes, which are clearly evident in Figure 5b. The large intercept and near-zero slope parameter during the third regime should be, therefore, less surprising. Malaysian slope parameters are shallower (particularly in the initial South Korea–Malaysia regime with a slope of 0.065), but it is not clear why this is so, nor is the reason for the revisions obvious since there are no spikes in spot imports during the last regimes. It is worth noting both that the length of the price series and minimum segment sizes are relatively short and that a number of the estimated oil price weights lie relatively far from the $[0, 1]$ interval. In these cases, the model may be overfitting the data with long lags.

Spanish pricing formulas appear to have much shallower slope parameters in the range 0.68–0.90, which is roughly what the tails of Japanese S-curves specify. Because of this, LNG prices have increased much less in Spain than for Asian buyers, even though Spain competes with Asia for Qatari cargoes. This is likely because Spain also has access to European gas markets and, therefore, more elastic demand for LNG. Furthermore, transportation costs may be prohibitively high for African LNG to reach Asia (Africa–Asia trade has not historically been large), so African nations may not be able to arbitrage away these differences. These large differences in pricing formulas suggests that the Atlantic and Pacific markets for LNG are still weakly arbitraged. Further work is needed to determine whether Spanish LNG prices are more tightly linked to continental European gas prices or oil markets.

7 Conclusion

This paper systematically examines the empirical relationship between LNG and crude oil prices using customs data from Japan, South Korea, Taiwan and Spain. Historically, these are the four most significant players in the LNG market, and the price series I examine correspond to a large portion of these countries' imports. Using techniques for detecting multiple unknown structural breaks in cointegrated regressions, I estimate the number and location of breaks in each relationship, as well as the implied LNG pricing formulas. Previous literature has focused on convergence of regional prices, ignoring the underlying price-setting mechanisms in LTCs. I take a complementary approach, ignoring statistical measures of price convergence but estimating the actual pricing terms.

Japanese contracts appear to be revised more often than those of other countries. Some of the revisions correspond to the use of S-curves in pricing formulas, and some, to the elimination of S-curves in contracts when markets were tighter and LNG was trading at a relatively steep discount to crude oil on an energy-parity basis. The fact that pricing formulas did not change much in response to the Fukushima incident is consistent with the idea that long-term LNG contracts function as a form of insurance against unexpected shocks, stabilizing cash flows and allowing firms to finance investment projects with more debt.

Unfortunately, it is unclear to what degree estimated structural breaks correspond to the presence of multiple SPAs that underlie Japanese imports, though for some relationships (such as the one with Brunei), there is only one contract. Korean and Taiwanese contracts, which are all linked to a single national importer in each country, are much simpler and have undergone little revision, if any. Additionally, slope-intercept formulas were historically steeper, meaning these two countries felt the impact of higher oil prices more than Japan did. In recent years, however, the slopes of Japanese contracts have resembled those in South Korea and Taiwan more. Spanish contracts are substantially different from the Asian contracts, with few revisions and much shallower slopes.

The focus of this paper has been on how and when empirical pricing relationships in LNG markets have changed as a way to understand how underlying pricing formulas in long-term contracts have developed. It appears that LNG prices continue to be indexed to oil prices, though the relationship between prices can be substantially more complex than industry rules of thumb predict.

Importer	Exporter	Start	End	T	Mean	Variance	D-Fuller	SSR_{Brent}	SSR_{JCC}	Benchmark	Lag
Spain	Algeria	Jan 1995	May 2014	233	5.731	10.779	-0.048	96.146	98.775	Brent	4
Spain	Nigeria	Feb 2000	May 2014	172	6.272	6.668	0.192	39.559	34.634	JCC	4
Spain	Qatar	Apr 2003	May 2014	134	7.421	6.049	0.097	47.723	45.390	JCC	2
Japan	UAE	Jun 1993	Jun 2014	253	7.293	24.289	1.271	414.702	361.030	JCC	3
Japan	Australia	Aug 1989	Jun 2014	299	6.620	18.310	0.996	316.334	279.197	JCC	3
Japan	Brunei	Jan 1988	Jun 2014	318	6.567	23.707	1.516	523.097	423.174	JCC	4
Japan	Indonesia	Jan 1988	Jun 2014	318	6.779	22.029	1.675	444.058	430.300	JCC	0
Japan	Malaysia	Jan 1988	Jun 2014	318	6.683	23.726	1.406	363.670	310.973	JCC	3
Japan	Qatar	Jan 1997	Jun 2014	210	8.571	26.223	1.222	248.896	185.442	JCC	3
Japan	USA	Jan 1988	Mar 2011	279	4.793	5.596	2.686	415.796	402.430	JCC	3
South Korea	Indonesia	Jul 2001	Jun 2014	156	10.246	15.625	-0.072	309.864	319.968	Brent	1
South Korea	Malaysia	Sep 2003	Jun 2014	130	8.771	13.201	-0.585	626.855	589.964	JCC	2
South Korea	Oman	Apr 2000	Jun 2014	171	10.868	28.585	0.577	55.020	49.365	JCC	3
South Korea	Qatar	Sep 1999	Jun 2014	178	10.526	28.282	0.988	50.432	31.512	JCC	4
Taiwan	Indonesia	Jan 2000	May 2014	173	11.152	28.503	0.297	52.329	65.711	Brent	1
Taiwan	Malaysia	Feb 2005	May 2014	112	12.043	28.265	-1.026	733.081	664.606	JCC	2
Crude	JCC	Jan 1988	Jun 2014	318	44.471	1223.085	0.916				
Crude	Brent	May 1987	Aug 2014	328	44.166	1211.863	0.492				

"D-Fuller" is the Dickey-Fuller test for a unit root with no trend or drift, and it is not rejected at the 10% level for any series.

"Lag" is the lag of the oil benchmark in the previous column which has the highest correlation with the series.

SSR_i is the sum of squared residuals from the following regression model: $P_{LNG,t} = \alpha + \sum_{k=0}^{p} \beta_i P_{i,t-k} + u_t$. If $SSR_i < SSR_j$, the benchmark is i.

Table 1: Summary Statistics

Japan–Australia: $T = 299$ $\epsilon T = 45$ $\epsilon = 0.151$

					$Break_{k+1}$
$k = 1$				2008(12)	2005(3)
$k = 2$			2005(3)	2008(12)	2000(3)
$k = 3$		2000(3)	2005(3)	2008(12)	1995(9)
$k = 4$	1995(9)	2000(5)	2005(3)	2008(12)	—
	$k = 0$	$k = 1$	$k = 2$	$k = 3$	$k = 4$
BIC	0.06	-1.37	-1.51	-1.45	-1.34
LWZ	0.20	-1.07	-1.05	-0.82	-0.56
SSR_k		59.38	44.92	41.97	40.83
F_T^*		21.23	14.47	9.94	7.53
$F_T^*(k+1\|k)$		14.72	5.76	2.38	-0.00
$\min(\Delta T_i)$		66	45	45	45

Japan–Brunei: $T = 311$ $\epsilon T = 47$ $\epsilon = 0.151$

					$Break_{k+1}$
$k = 1$				2007(12)	2003(12)
$k = 2$			2004(9)	2008(8)	2000(3)
$k = 3$		2000(3)	2004(9)	2008(8)	1993(1)
$k = 4$	1993(1)	2000(4)	2004(9)	2008(8)	—
	$k = 0$	$k = 1$	$k = 2$	$k = 3$	$k = 4$
BIC	0.44	-0.75	-0.94	-0.78	-0.60
LWZ	0.64	-0.32	-0.28	0.11	0.51
SSR_k		103.18	71.07	69.43	69.04
F_T^*		15.04	9.72	6.31	4.60
$F_T^*(k+1\|k)$		7.68	2.72	0.64	-0.00
$\min(\Delta T_i)$		78	47	47	47

Japan–Indonesia: $T = 315$ $\epsilon T = 36$ $\epsilon = 0.114$

					$Break_{k+1}$
$k = 1$				2011(3)	2004(12)
$k = 2$			2004(12)	2011(3)	2001(8)
$k = 3$		2001(8)	2004(12)	2011(3)	1997(5)
$k = 4$	1997(5)	2001(8)	2004(12)	2011(3)	2007(12)
	$k = 0$	$k = 1$	$k = 2$	$k = 3$	$k = 4$
BIC	0.42	-1.97	-2.57	-2.55	-2.49
LWZ	0.53	-1.73	-2.19	-2.03	-1.84
SSR_k		35.79	17.61	16.22	15.38
F_T^*		22.30	22.63	16.65	13.24
$F_T^*(k+1\|k)$		37.26	10.64	8.23	8.31
$\min(\Delta T_i)$		39	39	39	39

Japan–Malaysia: $T = 314$ $\epsilon T = 48$ $\epsilon = 0.153$

					$Break_{k+1}$
$k = 1$				2008(9)	2004(9)
$k = 2$			2004(9)	2008(9)	2000(2)
$k = 3$		2000(2)	2005(2)	2009(2)	1996(2)
$k = 4$	1996(6)	2000(6)	2005(2)	2009(2)	1992(4)
	$k = 0$	$k = 1$	$k = 2$	$k = 3$	$k = 4$
BIC	0.12	-1.39	-1.71	-1.66	-1.54
LWZ	0.26	-1.09	-1.26	-1.05	-0.77
SSR_k		61.91	39.32	36.50	36.20
F_T^*		21.52	17.38	6.07	4.49
$F_T^*(k+1\|k)$		27.04	6.19	0.52	0.14
$\min(\Delta T_i)$		69	48	48	48

Japan–Qatar: $T = 210$ $\epsilon T = 32$ $\epsilon = 0.152$

					$Break_{k+1}$
$k = 1$			2008(8)		2011(4)
$k = 2$		2005(9)	2009(4)		2000(8)
$k = 3$		2005(1)	2008(8)	2011(4)	2000(8)
$k = 4$	2000(8)	2005(12)	2008(8)	2011(4)	2003(4)
	$k = 0$	$k = 1$	$k = 2$	$k = 3$	$k = 4$
BIC	0.08	-1.29	-1.42	-1.42	-1.34
LWZ	0.30	-0.82	-0.70	-0.45	-0.11
SSR_k		37.60	26.25	20.73	17.88
F_T^*		19.94	14.08	10.74	9.36
$F_T^*(k+1\|k)$		20.79	14.06	13.13	1.87
$\min(\Delta T_i)$		70	43	32	32

Japan–UAE: $T = 253$ $\epsilon T = 38$ $\epsilon = 0.15$

					$Break_{k+1}$
$k = 1$				2009(2)	2004(12)
$k = 2$			2004(12)	2009(2)	1996(10)
$k = 3$		2000(4)	2005(10)	2009(2)	1996(10)
$k = 4$	1996(10)	2000(5)	2005(10)	2009(2)	—
	$k = 0$	$k = 1$	$k = 2$	$k = 3$	$k = 4$
BIC	0.51	-1.60	-1.79	-1.84	-1.87
LWZ	0.66	-1.27	-1.29	-1.16	-1.01
SSR_k		38.50	27.19	22.19	18.60
F_T^*		18.53	13.21	10.19	9.19
$F_T^*(k+1\|k)$		14.40	12.85	14.97	0.00
$\min(\Delta T_i)$		64	50	40	40

Japan–USA: $T = 276$ $\epsilon T = 24$ $\epsilon = 0.087$

					$Break_{k+1}$
$k = 1$				2009(3)	2004(8)
$k = 2$			2004(8)	2009(3)	1999(10)
$k = 3$		1999(10)	2005(12)	2009(3)	1991(3)
$k = 4$	1991(3)	1999(11)	2005(12)	2009(3)	1995(9)
	$k = 0$	$k = 1$	$k = 2$	$k = 3$	$k = 4$
BIC	0.49	-1.80	-3.01	-3.25	-3.43
LWZ	0.61	-1.53	-2.60	-2.70	-2.73
SSR_k		36.64	9.64	6.67	4.97
F_T^*		16.40	15.39	13.26	12.55
$F_T^*(k+1\|k)$		27.98	18.10	30.81	12.75
$\min(\Delta T_i)$		24	24	24	24

Table 2: Break tests (Japan)

South Korea–Indonesia: $T = 156 \quad \epsilon T = 24 \quad \epsilon = 0.154$

					$Break_{k+1}$
$k = 1$			2009(1)		2011(7)
$k = 2$			2008(12)	2011(7)	2005(4)
$k = 3$		2007(6)	2009(6)	2011(7)	2005(6)
$k = 4$	2005(6)	2007(6)	2009(6)	2011(7)	2003(6)
	$k = 0$	$k = 1$	$k = 2$	$k = 3$	$k = 4$
BIC	0.86	-0.41	-0.56	-0.46	-0.33
LWZ	0.96	-0.19	-0.21	0.02	0.27
SSR_k		82.73	62.28	60.61	60.46
F_T^*		75.03	28.10	17.89	13.09
$F_T^*(k+1\|k)$		34.28	2.07	0.48	0.12
$\min(\Delta T_i)$		65	31	24	24

South Korea–Malaysia: $T = 130 \quad \epsilon T = 20 \quad \epsilon = 0.154$

					$Break_{k+1}$
$k = 1$				2012(4)	2008(12)
$k = 2$		2008(12)		2012(4)	2010(8)
$k = 3$		2008(12)	2011(2)	2012(10)	2006(11)
$k = 4$	2006(11)	2008(12)	2011(2)	2012(10)	—
	$k = 0$	$k = 1$	$k = 2$	$k = 3$	$k = 4$
BIC	1.77	1.48	1.57	1.65	1.83
LWZ	1.90	1.79	2.05	2.31	2.67
SSR_k		408.54	369.99	332.46	330.71
F_T^*		12.90	7.47	5.25	3.82
$F_T^*(k+1\|k)$		3.93	1.80	0.22	-0.00
$\min(\Delta T_i)$		26	26	20	20

South Korea–Oman: $T = 171 \quad \epsilon T = 26 \quad \epsilon = 0.152$

					$Break_{k+1}$
$k = 1$				2010(8)	2008(6)
$k = 2$			2008(6)	2010(8)	2006(2)
$k = 3$		2005(6)	2007(8)	2010(8)	2003(3)
$k = 4$	2003(11)	2006(2)	2008(6)	2010(8)	—
	$k = 0$	$k = 1$	$k = 2$	$k = 3$	$k = 4$
BIC	-1.03	-0.94	-0.96	-0.92	-0.79
LWZ	-0.85	-0.54	-0.35	-0.09	0.26
SSR_k		45.37	35.75	30.24	27.88
F_T^*		8.05	10.48	10.16	7.74
$F_T^*(k+1\|k)$		18.21	10.92	3.21	0.00
$\min(\Delta T_i)$		46	26	26	26

South Korea–Qatar: $T = 178 \quad \epsilon T = 27 \quad \epsilon = 0.152$

					$Break_{k+1}$
$k = 1$			2009(1)		2011(8)
$k = 2$			2008(3)	2011(8)	2005(12)
$k = 3$		2006(1)	2008(4)	2011(8)	2003(7)
$k = 4$	2003(7)	2006(2)	2008(5)	2011(8)	—
	$k = 0$	$k = 1$	$k = 2$	$k = 3$	$k = 4$
BIC	-1.50	-1.39	-1.28	-1.10	-0.91
LWZ	-1.26	-0.88	-0.50	-0.03	0.44
SSR_k		26.97	23.22	21.49	20.00
F_T^*		10.80	8.76	6.34	5.18
$F_T^*(k+1\|k)$		9.13	6.24	5.78	-0.00
$\min(\Delta T_i)$		65	34	27	27

Table 3: Break tests (South Korea)

Spain–Algeria: $T = 233 \quad \epsilon T = 35 \quad \epsilon = 0.15$

					$Break_{k+1}$
$k=1$			2007(1)		2011(6)
$k=2$			2007(1)	2011(6)	2000(5)
$k=3$		2005(1)	2007(12)	2011(6)	2000(5)
$k=4$	2000(5)	2005(1)	2007(12)	2011(6)	—
	$k=0$	$k=1$	$k=2$	$k=3$	$k=4$
BIC	-0.47	-0.55	-0.59	-0.55	-0.43
LWZ	-0.32	-0.21	-0.07	0.16	0.47
SSR_k		99.01	80.64	71.26	68.40
F_T^*		19.72	16.51	15.09	12.00
$F_T^*(k+1\|k)$		22.67	5.57	6.64	0.00
$\min(\Delta T_i)$		88	35	35	35

Spain–Nigeria: $T = 172 \quad \epsilon T = 26 \quad \epsilon = 0.151$

					$Break_{k+1}$
$k=1$			2008(12)		2002(3)
$k=2$			2008(9)	2010(11)	2002(3)
$k=3$	2002(3)		2008(9)	2010(11)	2005(11)
$k=4$	2002(3)	2005(11)	2008(9)	2010(11)	—
	$k=0$	$k=1$	$k=2$	$k=3$	$k=4$
BIC	-1.28	-1.30	-1.16	-1.00	-0.82
LWZ	-1.04	-0.79	-0.36	0.08	0.55
SSR_k		28.09	24.83	22.16	20.30
F_T^*		22.16	15.69	14.67	15.06
$F_T^*(k+1\|k)$		11.72	15.63	14.45	0.00
$\min(\Delta T_i)$		65	26	26	26

Spain–Qatar: $T = 134 \quad \epsilon T = 21 \quad \epsilon = 0.157$

					$Break_{k+1}$
$k=1$		2008(12)			2012(8)
$k=2$		2008(12)		2012(8)	2010(9)
$k=3$		2008(12)	2010(9)	2012(8)	2006(2)
$k=4$	2006(2)	2008(12)	2010(9)	2012(8)	—
	$k=0$	$k=1$	$k=2$	$k=3$	$k=4$
BIC	-0.78	-0.69	-0.62	-0.36	-0.08
LWZ	-0.51	-0.11	0.27	0.86	1.47
SSR_k		36.07	27.78	25.99	24.77
F_T^*		10.70	10.73	7.65	5.88
$F_T^*(k+1\|k)$		14.96	5.68	4.49	-0.00
$\min(\Delta T_i)$		65	21	21	21

Taiwan–Indonesia: $T = 173 \quad \epsilon T = 26 \quad \epsilon = 0.15$

					$Break_{k+1}$
$k=1$			2008(10)		2011(4)
$k=2$			2008(10)	2011(4)	2006(8)
$k=3$		2006(8)	2008(10)	2011(4)	2002(10)
$k=4$	2002(10)	2006(8)	2008(10)	2011(4)	—
	$k=0$	$k=1$	$k=2$	$k=3$	$k=4$
BIC	-0.99	-0.99	-0.89	-0.79	-0.67
LWZ	-0.90	-0.78	-0.56	-0.33	-0.09
SSR_k		52.04	51.10	50.35	50.20
F_T^*		17.30	9.59	6.98	5.22
$F_T^*(k+1\|k)$		2.86	2.37	0.47	0.00
$\min(\Delta T_i)$		67	30	26	26

Taiwan–Malaysia: $T = 112 \quad \epsilon T = 28 \quad \epsilon = 0.25$

					$Break_{k+1}$
$k=1$	2008(12)				2011(11)
$k=2$	2008(11)	2011(6)			—
	$k=0$	$k=1$	$k=2$	$k=3$	$k=4$
BIC	2.06	1.57	1.59		
LWZ	2.28	2.05	2.34		
SSR_k		311.35	237.20		
F_T^*		49.37	21.70		
$F_T^*(k+1\|k)$		6.93	-0.00		
$\min(\Delta T_i)$		47	31		

Table 4: Break tests (Spain and Taiwan)

	Japan–Australia			Japan–Brunei		Japan–Indonesia			Japan–Malaysia		
Start	1989(08)	2005(04)	2009(01)	1988(08)	2008(01)	1988(04)	2005(01)	2011(04)	1988(05)	2004(10)	2008(10)
End	2005(03)	2008(12)	2014(06)	2007(12)	2014(06)	2004(12)	2011(03)	2014(06)	2004(09)	2008(09)	2014(06)
c	1.869***	0.078	1.218	2.309***	−0.429	0.951***	2.337***	−3.352	1.247***	0.674	−1.156*
	(0.128)	(0.533)	(0.656)	(0.149)	(1.502)	(0.118)	(0.176)	(2.483)	(0.098)	(0.715)	(0.438)
δ	0.095***	0.108***	0.129***	0.064***	0.156***	0.141***	0.089***	0.188***	0.120***	0.101***	0.166***
	(0.006)	(0.007)	(0.007)	(0.005)	(0.015)	(0.005)	(0.002)	(0.022)	(0.005)	(0.012)	(0.004)
θ_0	−0.189	0.095	0.041	0.296	0.055	0.993***	0.932***	0.342**	0.050	0.439	0.338**
	(0.256)	(0.257)	(0.238)	(0.623)	(0.337)	(0.160)	(0.093)	(0.106)	(0.150)	(0.474)	(0.101)
θ_1	−0.034	−0.324	0.084	−0.145	0.134	0.075	0.147	0.013	−0.102	0.070	−0.111
	(0.453)	(0.505)	(0.373)	(1.066)	(0.612)	(0.288)	(0.175)	(0.169)	(0.277)	(0.845)	(0.181)
θ_2	0.242	0.375	0.107	−0.283	0.044	−0.085	−0.092	−0.094	0.161	0.093	−0.056
	(0.476)	(0.592)	(0.364)	(1.108)	(0.612)	(0.293)	(0.174)	(0.168)	(0.296)	(0.910)	(0.178)
θ_3	0.845	0.816	0.625	0.409	−0.014	0.017	0.013	0.739***	−0.014	0.711	0.466*
	(0.465)	(0.621)	(0.340)	(1.109)	(0.632)	(0.168)	(0.099)	(0.108)	(0.279)	(0.920)	(0.178)
θ_4	0.136	0.037	0.142	0.655	0.518				0.905***	−0.313	0.363***
	(0.262)	(0.377)	(0.186)	(1.140)	(0.627)				(0.152)	(0.630)	(0.096)
θ_5				0.160	−0.171						
				(1.159)	(0.603)						
θ_6				0.689	−0.094						
				(1.123)	(0.604)						
θ_7				−0.781	0.528						
				(0.685)	(0.327)						
T	188	45	66	233	78	201	75	39	197	48	69
$\hat{\sigma}_i$	0.174	0.576	0.664	0.338	1.060	0.152	0.287	0.463	0.161	0.607	0.548
$\hat{\sigma}_{LR,i}$	0.469	0.922	1.086	1.063	2.413	0.440	0.376	0.499	0.335	1.277	0.735
$\hat{\rho}_i$	0.797	0.538	0.467	0.801	0.757	0.812	0.363	0.217	0.690	0.662	0.445
DW_i	0.370	0.926	1.065	0.361	0.501	0.387	1.210	1.520	0.613	0.587	1.090
$\Pr(\mathrm{DW}_i)$	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.047	0.000	0.000	0.000
$\hat{\sigma}$		0.400		0.593			0.242			0.364	
$\hat{\sigma}_{LR}$		0.720		1.550			0.372			0.611	
$\hat{\rho}$		0.511		0.736			0.384			0.452	
D-Fuller		-9.778		-6.882			-11.810			-10.862	

$^{***}p < 0.001$, $^{**}p < 0.01$, $^{*}p < 0.05$ Standard errors in parenthesis are calculated by the delta-method. DW is the Durbin-Watson test for first-order autocorrelation.

Table 5: Models I

	Japan–Qatar			Japan–UAE			Japan–USA				
Start	1997(01)	2005(10)	2009(05)	1993(06)	2005(01)	2009(03)	1988(04)	1991(04)	1999(12)	2006(01)	2009(04)
End	2005(09)	2009(04)	2014(06)	2004(12)	2009(02)	2014(06)	1991(03)	1999(11)	2005(12)	2009(03)	2011(03)
c	1.890^{***}	-1.848^{**}	2.305^{***}	1.335^{***}	0.485	-0.389	1.452^{***}	0.974^{***}	2.470^{***}	3.447^{***}	0.823^{**}
	(0.211)	(0.565)	(0.308)	(0.251)	(0.321)	(0.319)	(0.131)	(0.173)	(0.059)	(0.059)	(0.236)
δ	0.095^{***}	0.156^{***}	0.132^{***}	0.113^{***}	0.106^{***}	0.155^{***}	0.096^{***}	0.134^{***}	0.070^{***}	0.041^{***}	0.146^{***}
	(0.008)	(0.007)	(0.003)	(0.011)	(0.004)	(0.003)	(0.007)	(0.010)	(0.002)	(0.001)	(0.003)
θ_0	-0.094	-0.027	0.029	0.072	-0.116	0.173	0.080	0.183	0.076	0.042	-0.015
	(0.392)	(0.161)	(0.097)	(0.442)	(0.177)	(0.095)	(0.213)	(0.226)	(0.118)	(0.062)	(0.063)
θ_1	0.038	0.103	-0.185	-0.121	-0.104	-0.242	0.287	-0.054	-0.060	-0.493^{***}	-0.043
	(0.642)	(0.310)	(0.144)	(0.720)	(0.338)	(0.145)	(0.474)	(0.389)	(0.188)	(0.116)	(0.086)
θ_2	0.350	0.172	0.422^{**}	0.206	0.292	0.226	-1.038^{*}	-0.746^{*}	0.086	0.576^{***}	0.088
	(0.645)	(0.320)	(0.146)	(0.757)	(0.342)	(0.144)	(0.486)	(0.372)	(0.189)	(0.120)	(0.087)
θ_3	0.267	0.116	0.452^{**}	-0.224	0.444	0.630^{***}	1.671^{***}	1.617^{***}	0.899^{***}	0.875^{***}	0.970^{***}
	(0.645)	(0.336)	(0.144)	(0.754)	(0.344)	(0.142)	(0.223)	(0.205)	(0.122)	(0.058)	(0.065)
θ_4	0.214	0.213	0.170	1.066^{*}	0.484^{*}	0.212^{*}					
	(0.669)	(0.333)	(0.142)	(0.460)	(0.184)	(0.081)					
θ_5	0.110	-0.147	0.134								
	(0.676)	(0.322)	(0.139)								
θ_6	0.116	0.570^{**}	-0.021								
	(0.426)	(0.171)	(0.080)								
T	105	43	62	139	50	64	36	104	73	39	24
$\hat{\sigma}_i$	0.216	0.566	0.441	0.279	0.432	0.386	0.126	0.163	0.084	0.103	0.230
$\hat{\sigma}_{LR,i}$	0.625	0.812	0.416	0.788	0.619	0.504	0.157	0.247	0.133	0.085	0.179
$\hat{\rho}_i$	0.886	0.502	0.043	0.791	0.376	0.318	0.403	0.493	0.462	-0.251	-0.228
DW_i	0.205	0.971	1.889	0.405	1.247	1.359	1.137	1.002	1.051	2.455	2.434
$\Pr(\mathrm{DW}_i)$	0.000	0.000	0.271	0.000	0.001	0.003	0.001	0.000	0.000	0.870	0.795
$\hat{\sigma}$		0.376			0.340				0.141		
$\hat{\sigma}_{LR}$		0.540			0.627				0.180		
$\hat{\rho}$		0.391			0.511				0.284		
D-Fuller		-9.464			-8.983				-12.341		

$^{***}p < 0.001$, $^{**}p < 0.01$, $^{*}p < 0.05$ Standard errors in parenthesis are calculated by the delta-method. DW is the Durbin-Watson test for first-order autocorrelation.

Table 6: Models II

	South Korea–Indonesia			South Korea–Malaysia		South Korea–Oman	South Korea–Qatar
Start	2001(07)	2009(01)	2011(08)	2003(09)	2012(05)	2000(04)	1999(09)
End	2008(12)	2011(07)	2014(06)	2012(04)	2014(06)	2014(06)	2014(06)
c	0.738^{***}	3.951^{***}	13.192^{***}	2.958^{***}	-36.246	0.719^{***}	0.817^{***}
	(0.069)	(0.625)	(2.794)	(0.409)	(45.232)	(0.139)	(0.098)
δ	0.157^{***}	0.077^{***}	0.006	0.065^{***}	0.452	0.156^{***}	0.154^{***}
	(0.001)	(0.008)	(0.025)	(0.005)	(0.408)	(0.002)	(0.001)
θ_0	-0.086^{*}	0.309	-11.460	0.466	0.568	-0.082	-0.078
	(0.033)	(0.423)	(53.175)	(0.466)	(0.762)	(0.097)	(0.072)
θ_1	1.086^{***}	0.691	12.460	-1.412	-0.317	0.020	0.077
	(0.033)	(0.423)	(53.175)	(0.831)	(1.045)	(0.175)	(0.131)
θ_2				1.946^{***}	0.749	0.072	0.092
				(0.474)	(0.628)	(0.177)	(0.134)
θ_3						0.583^{**}	0.080
						(0.175)	(0.135)
θ_4						0.407^{***}	0.374^{**}
						(0.096)	(0.134)
θ_5							0.334^{*}
							(0.131)
θ_6							0.122
							(0.072)
T	90	31	35	104	26	171	178
$\hat{\sigma}_i$	0.310	1.001	0.899	1.281	3.333	0.557	0.431
$\hat{\sigma}_{LR,i}$	0.303	0.889	0.703	1.472	6.769	0.816	0.604
$\hat{\rho}_i$	-0.027	-0.095	-0.263	0.181	0.658	0.413	0.552
DW_i	2.049	2.181	2.472	1.639	0.627	1.171	0.891
$\Pr(\mathrm{DW}_i)$	0.526	0.624	0.891	0.022	0.000	0.000	0.000
$\hat{\sigma}$		0.651		1.830		0.557	0.431
$\hat{\sigma}_{LR}$		0.579		3.190		0.816	0.604
$\hat{\rho}$		-0.149		0.466		0.413	0.552
D-Fuller		-14.275		-6.602		-8.332	-7.041

$^{***}p < 0.001$, $^{**}p < 0.01$, $^{*}p < 0.05$ Standard errors in parenthesis are calculated by the delta-method.

DW is the Durbin-Watson test for first-order autocorrelation.

Table 7: Models III

	Spain–Algeria			Spain–Nigeria		Spain–Qatar	Taiwan–Indonesia		Taiwan–Malaysia	
Start	1995(01)	2007(02)	2011(07)	2000(02)	2009(01)	2003(04)	2000(01)	2008(11)	2005(02)	2009(01)
End	2007(01)	2011(06)	2014(05)	2008(12)	2014(05)	2014(05)	2008(10)	2014(05)	2008(12)	2014(05)
c	0.942^{***}	2.859^{**}	12.409^{**}	1.082^{***}	2.109^{***}	1.360^{***}	1.008^{***}	0.907^{**}	-0.260	-1.673
	(0.098)	(0.993)	(4.281)	(0.140)	(0.306)	(0.215)	(0.118)	(0.287)	(0.981)	(0.960)
δ	0.090^{***}	0.073^{***}	-0.014	0.083^{***}	0.068^{***}	0.080^{***}	0.152^{***}	0.158^{***}	0.131^{***}	0.173^{***}
	(0.003)	(0.012)	(0.039)	(0.003)	(0.003)	(0.003)	(0.002)	(0.003)	(0.014)	(0.010)
θ_0	-0.034	0.375	3.949	0.419^{*}	0.889^{***}	0.179	0.162^{*}	-0.130	0.524	0.268
	(0.184)	(0.449)	(9.938)	(0.202)	(0.191)	(0.208)	(0.070)	(0.072)	(0.397)	(0.257)
θ_1	0.078	0.084	0.603	-0.466	-0.710^{*}	-0.218	0.838^{***}	1.130^{***}	-2.802^{**}	-0.327
	(0.267)	(0.765)	(3.074)	(0.390)	(0.310)	(0.377)	(0.070)	(0.072)	(0.843)	(0.406)
θ_2	-0.003	-0.654	-7.387	-0.012	0.086	0.521			2.194^{*}	0.797
	(0.274)	(0.729)	(21.007)	(0.440)	(0.309)	(0.387)			(0.994)	(0.402)
θ_3	0.184	-0.226	6.266	0.314	0.317	0.112			-0.842	-0.019
	(0.276)	(0.783)	(17.135)	(0.465)	(0.292)	(0.391)			(0.961)	(0.367)
θ_4	0.775^{***}	1.420^{**}	-2.431	0.266	0.259	0.087			1.927^{**}	0.281
	(0.192)	(0.446)	(7.633)	(0.472)	(0.288)	(0.386)			(0.556)	(0.201)
θ_5				0.108	-0.273	0.008				
				(0.468)	(0.281)	(0.378)				
θ_6				0.371	0.434^{**}	0.311				
				(0.304)	(0.152)	(0.207)				
T	145	53	35	107	65	134	106	67	47	65
$\hat{\sigma}_i$	0.331	0.977	0.842	0.367	0.509	0.602	0.556	0.561	1.985	1.594
$\hat{\sigma}_{LR,i}$	0.549	1.492	0.664	0.627	0.471	0.883	0.588	0.546	1.783	1.587
$\hat{\rho}_i$	0.413	0.518	-0.236	0.549	-0.059	0.382	0.113	-0.031	-0.084	0.094
DW_i	1.165	0.961	2.455	0.899	2.110	1.224	1.771	1.768	1.849	1.805
$\Pr(\mathrm{DW}_i)$	0.000	0.000	0.869	0.000	0.592	0.000	0.091	0.131	0.223	0.168
$\hat{\sigma}$	0.612			0.424		0.602	0.558		1.765	
$\hat{\sigma}_{LR}$	0.828			0.519		0.883	0.570		1.678	
$\hat{\rho}$	0.307			0.222		0.382	0.056		0.024	
D-Fuller	-10.983			-10.428		-7.527	-12.292		-10.187	

$^{***}p < 0.001$, $^{**}p < 0.01$, $^{*}p < 0.05$ Standard errors in parenthesis are calculated by the delta-method. DW is the Durbin-Watson test for first-order autocorrelation.

Table 8: Models IV

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A Estimation of structural breaks

Estimation and testing of multiple, unknown change points in a stationary context was addressed by Bai and Perron (1998, 2003), while Kejriwal and Perron (2008b, 2010) extended the tests to a cointegrated context. Let $\boldsymbol{c}$ and $\boldsymbol{\delta}$ be the m+1-length vectors of regime-specific coefficients c_j and δ_j. Similarly, let $\boldsymbol{\pi}_j$ be the $p+1$ vector of coefficients $\pi_0, \ldots, \pi_p$ and $\boldsymbol{\pi}$ be the $(m+1) \times (p+1)$ vector of $\boldsymbol{\pi}_j$ vectors. For each m-partition $(T_1, \ldots, T_m)$, OLS estimates $\hat{\boldsymbol{c}}(\{T_j\})$, $\hat{\boldsymbol{\delta}}(\{T_j\})$ and $\hat{\boldsymbol{\pi}}(\{T_j\})$ minimize

$$SSR_T(T_1, \ldots, T_m) = \min_{c_j, \delta_j, \boldsymbol{\pi}_j} \sum_{j=1}^{m+1} \sum_{t=T_{j-1}+1}^{T_j} [y_t - c_j - z_t \delta_j - \Delta \boldsymbol{z}_t' \boldsymbol{\pi}_j]^2.$$

Define $S_T(T_1, \ldots, T_m)$ as the value of SSR_T when the estimates are substituted into SSR_T. Then the estimated break-points are just

$$(\hat{T}_1, \ldots, \hat{T}_m) = \underset{T_1, \ldots, T_m}{\arg\min}\, S_T(T_1, \ldots, T_m)$$

where segments the length of each segment $(\Delta T_j = T_j - T_{j-1})$ is at least ϵT: $\Delta T_j \geq \epsilon T$ and $\epsilon > 0$. Under assumptions A1–A8 of Kejriwal and Perron (2008b), the vector of estimated break-fractions converges to the true value: $\hat{\lambda} = (\hat{T}_1/T, \ldots, \hat{T}_m/T) \xrightarrow{p} \lambda^0$. The convergence of coefficient estimates immediately follows. Note that assumption A1 does *not* require that z have the same distribution over all segments, and A4 allows for conditional heteroskedasticity and autocorrelation in u_t. This means that breakpoints are consistently estimated *even if the variance of u_t changes between regimes.*

A.1 Testing break candidates

There are three tests of interest to determine whether structural breaks are present and, if so, how many are present. Kejriwal and Perron (2010) provide critical values for the tests when $\epsilon \in \{0.15, 0.20, 0.25\}$. Define the number of coefficients to be $q^* = 1 + 1 + (p+1)$ and SSR_k as the sum of OLS squared residuals when k breaks are allowed. The first test described in Kejriwal and Perron (2008b, 2010) is a Wald F-test for a fixed $m = k$ versus 0 breaks scaled by the number of coefficients allowed to change (q):

$$F_T(\lambda, k) = \frac{T - q^*(k+1)}{k} \times \frac{SSR_0 - SSR_k}{SSR_k}$$

The test of interest, the sup-Wald test, is simply $\sup_{\lambda \in \Lambda_\epsilon^k} F_T(\lambda, k)$. The maximizing break-vector of this test is equal to the minimizer of S_T. The test can be made robust to autocorrelation of u_t by a simple scaling

$$F_T^*(k) = (\hat{\sigma}_u^2 / \hat{\sigma}^2) \times F_T(k) \tag{3}$$

where $\hat{\sigma}_u^2 = SSR_k/T$ and $\hat{\sigma}^2$ is the estimator of the long-run variance (σ^2) proposed by Kejriwal and Perron (2010). Specifically, it is a modified Andrews (1991) estimator of the long-run variance computed as

$$\hat{\sigma}^2 = T^{-1}\sum_{t=1}^{T}\tilde{u}_t^2 + 2T^{-1}\sum_{j=1}^{T-1} w(j/\hat{h}) \sum_{t=j+1}^{T} \tilde{u}_t\tilde{u}_{t-j}. \tag{4}$$

The residuals $\tilde{u}_t$ are from the model under the null hypothesis (0 breaks for the sup Wald test, $F_T^*(k)$, or $k+1$ breaks for the sequential test, $F_T^*(k+1|k)$), but the bandwidth is estimated using the residuals under the alternative hypothesis of k (or $k+1$) breaks: $\hat{h} = 1.3221[\hat{\alpha}(2)T]^{1/5}$, $\hat{\alpha}(2) = 4\hat{\rho}^2/(1-\hat{\rho})^4$ and $\hat{\rho} = \sum_{t=2}^{T}\hat{u}_t\hat{u}_{t-1}/\sum_{t=2}^{T}\hat{u}_{t-1}^2$. Since $\hat{\lambda}$ is a superconsistent estimator for λ^0, we do not need to re-estimate the break-fractions with a robust estimator.[20] The second test is a double-maximum test for an unknown number of breaks $m = 1, \ldots, M$ versus 0 breaks:

$$UD\max F_T^*(M) = \max_{1\le m\le M}\ \sup_{\lambda\in\Lambda_\epsilon^m} F_T^*(\hat{\lambda}, m). \tag{5}$$

The final test is a sequential one for k versus $k+1$ breaks

$$F_T(k+1|k) = \max_{1\le j\le k+1}\ \sup_{\tau\in\Lambda_{j,\epsilon}} \frac{A_T(k) - B_T(\tau, k)}{SSR_{k+1}/T} \tag{6}$$

where $A_T(k) = SSR_T(\hat{T}_1, \ldots, \hat{T}_k)$ minimizes SSR with k breaks and the quantity $B(k,\tau) = SSR_T(\hat{T}_1, \ldots, \hat{T}_{j-1}, \tau, \hat{T}_j, \ldots, \hat{T}_k)$ is the minimum SSR when one break is allowed in addition to $\hat{T}_1, \ldots, \hat{T}_k$. Note that this break must be at least $\eta(T_j - T_{j-1})$ periods from the endpoints of any segment (the set $\lambda_{j,\eta} = \{\tau|\hat{T}_{j-1} + (\hat{T}_j - \hat{T}_{j-1})\eta \le \tau \le \hat{T}_j - (\hat{T}_j - \hat{T}_{j-1})\eta\}$). In earlier work, Kejriwal (2008) considers a robust version of the test similar to equation (3), which I use instead of (6):

$$F_T^*(k+1|k) = \max_{1\le j\le k+1}\ \sup_{\tau\in\Lambda_{j,\eta}} \frac{A_T(k) - B_T(\tau, k)}{\hat{\sigma}_{k+1}^2} \tag{7}$$

where $\hat{\sigma}_{k+1}^2$ is estimated as in equation (4).

Bai and Perron (1998, 2003) suggest first applying the $UD\max$ test for the presence of structural breaks, then using the sequential $F_T(k+1|k)$ test to determine the number of breaks (which corresponds to the point where the test fails to reject). This is the approach used by Kejriwal and Perron (2010). Bai and Perron (1998, 2003) and Kejriwal (2008) also mention that the BIC and a modified Schwarz criterion (LWZ) can be used to determine the number of breaks, though in Bai and Perron (2003), the authors note that the BIC and LWZ tend to overestimate the number of breaks when errors are serially correlated, as found by Perron (1997). These information criterion

[20] The tests assume that the error process does not change over time under the null hypothesis. In particular, the tests assume a constant variance, which seems unreasonable in this context. Incorporating changes in variance between regimes would require a modified set of tests. Given that the estimate of break-dates is super-consistent and valid even under regime-specific heteroskedasticity, it is unclear that modifications would add much insight. Nevertheless, because the variance of the error terms does not appear to be constant across regimes, I choose not to calculate confidence intervals for breakpoints since these, unlike breakpoint estimates, would not be consistent under regime-specific heteroskedasticity.

are calculated by Kejriwal (2008) as

$$BIC(m) = \log\left(\frac{SSR_m}{T}\right) + q^* \frac{\log(T)}{T} \tag{8}$$

$$LWZ(m) = \log\left(\frac{SSR_m}{T - q^*}\right) + \left(\frac{q^*}{T}\right) c_0 \log(T)^{2+\delta_0} \tag{9}$$

where $q^* = q(m+1) + m$ is the number of estimated parameters (q coefficients over $m+1$ regimes plus m breakpoints). I set $\delta_0 = 0.1$ and $c_0 = 0.299$ as suggested by Kejriwal (2008), who cites Liu et al. (1997).

B Extra plots

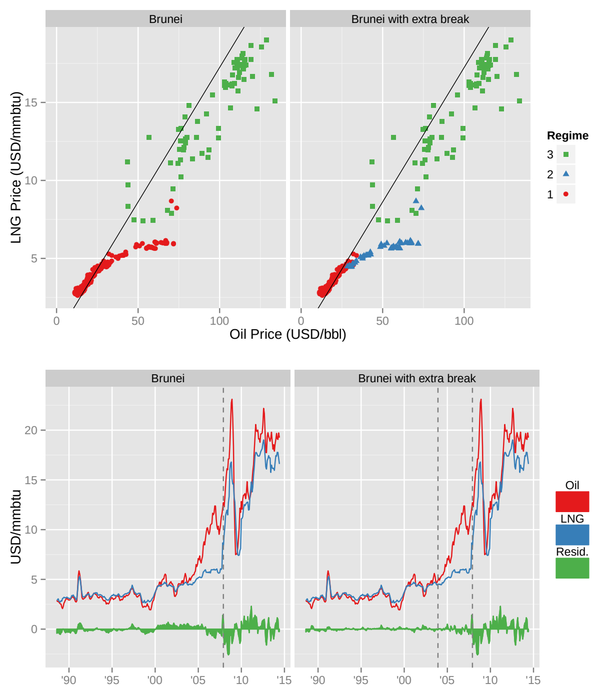

Figure 9: Japan–Brunei with extra break at 2003(12)