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Sanctions against Iran: An assessment of their global impact through the lens of international methanol prices ☆

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Abstract

Iran's energy and petrochemical exports have recently been restricted by a series of international sanctions. This paper focuses on one of the country's exports, namely methanol – a petrochemical increasingly used for fuel blending and traded at various locations worldwide – and empirically explores the relationships among the North American, European, and Asian markets to investigate the incidence of these sanctions. The analyses are conducted under a parity bounds framework based on Negassa and Myers (2007). The model was applied to the main methanol importing markets to estimate the effects of the sanctions on the degree of spatial integration. The findings document the occurrence of a complete reconfiguration of the spatial extent of the methanol markets. Under the sanctions, an increased degree of market integration was observed across the Atlantic, while fragmentation rose between Europe, South East Asia, and the two giant economies of China and India which both experienced lower prices.

Keywords: Iran, sanctions, law of one price, market integration, methanol.

JEL Classification: Q02, F51, F15, O13, Q48.

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1. Introduction

Over the last 35 years, the desire to set stringent limits on Iran's nuclear activities and the broader prospects for the accommodation of regime change have motivated a number of unilateral and multilateral sanctions against the Islamic Republic of Iran. Significant developments occurred after 2012 when the US and the European Union jointly adopted draconian measures aimed at restricting the country's ability to export oil, gas, and petrochemicals (Cordesman et al., 2014). By drastically prohibiting Iran's access to western-controlled shipping-related services (e.g., marine insurance, banking system) these economic sanctions *de facto* created an embargo. As some observers argued that this embargo was partly bypassed,[1] its effectiveness needs to be assessed. Surprisingly, despite the importance of these sanctions in foreign policy debates, their spatial incidence on the prices of energy-related commodities observed around the world has hitherto never been studied.

As this statement invites an empirical investigation, this paper examines the degree of integration and its evolution for methanol trade around the world following the sanctions. Methanol, a petrochemical product primarily produced from the methane component of natural gas and increasingly used for fuel blending, provides a relevant case study to assess the impacts of the sanctions for the following reasons. First, Iran is a major exporter of this basic petrochemical and this product dominates the country's petrochemical exports. Second, methanol is a globally traded commodity and there are active spot methanol markets at various locations worldwide, including Asia, Europe, and Northern America, which are supply–demand driven. Third, in contrast to most petroleum derivatives (e.g., refined products, naphtha), this is a globally standardized product. Therefore, its spatial price spreads are not caused by regional variations in quality standards. Altogether, these features make it possible to test whether the market definitions of Cournot and Marshall—that two regions are in the same economic market for a homogeneous good if the prices for that good differ by exactly the interregional transportation cost—are verified or not, and whether the sanctions have had an impact on market integration. Thus, in this paper we are particularly interested in identifying: (i) the geographic domain(s) over which the Law of One Price (LOOP) holds as an equilibrium condition after due allowance for arbitrage costs, (ii) the frequency at which that LOOP holds, and (iii) whether that frequency has changed following the sanctions.

The literature provides a large amount of empirical research which examines the degree of spatial integration between energy markets with the help of time-series techniques.[2] These studies typically

[1] *For example: "Iran finding some ways to evade sanctions, Treasury Department Says" (New York Times, January 10, 2013), "China floods Iran with cheap consumer goods in exchange for oil" (The Guardian, February 20, 2013).*

[2] *A tentative and non-exhaustive methodological clustering of these contributions includes: (i) early correlation-based analyses (Doane and Spulber, 1994); (ii) the use of Granger causality tests to examine the transmission of price shocks across markets (Doane and Spulber, 1994); (iii) the application of a co-integration test to investigate the existence of long-run common stochastic trends in the local prices as evidence of market integration using either a bivariate approach (De Vany and Walls, 1993; Serletis, 1997) or a multivariate one (Asche et al., 2002); (iv) the use of an autoregressive model of pairwise price differentials between geographically diverse locations to estimate the speeds of adjustment toward equilibrium (Cuddington and Wang, 2006); (v) the joint assessments of the degree of market integration and price transmission across markets using tests of co-integration and the corresponding error-correction models (Brown and Yücel, 2008).*

rely on local price data and assess the co-movements of prices at each market location. In these analyses, it is typically argued that high degrees of correlation and/or co-integration between the price series are evidence that the law of one price is being enforced through spatial arbitrages. These price-based empirical models usually provide useful insights into how local price shocks are transmitted to adjacent markets. However, these conventional approaches implicitly posit an ideal situation that overlooks some of the complexity associated with spatial arbitrage. For example, their application to a pair of markets exhibiting alternating periods of autarky and trade or fluctuating arbitrage costs could be a source of inferential danger (Baulch, 1997). To overcome these issues, researchers may consider more sophisticated techniques such as the Kalman filter approach to examine a time-varying degree of price convergence among spot markets (King and Cuc, 1996; Neumann et al., 2006) or a specification drawn from the wide class of nonlinear threshold autoregressive models (e.g., TAR models, Band-TAR models).

In this paper, we consider an empirical approach based on the parity bounds model (PBM) first introduced in Spiller and Huang (1986) to examine the extent of the integration of the wholesale gasoline markets in the Northeastern United States. Following an extension in Sexton et al. (1991), the PBM framework has been widely applied in agricultural economics (Baulch, 1997; Barrett et al., 2000; Barrett and Li, 2002; Park et al., 2002; Padilla-Bernal et al., 2003; Negassa and Myers, 2007, Moser et al., 2009). In contrast, and despite its origins, the PBM methodology is seldom found in the energy economics literature.[3] In a PBM, arbitrageurs are assumed to be profit-maximizing agents. Using that assumption, intermarket price spreads are examined using a "regime switching" specification which estimates the probability of observing each of a series of trade regimes. For example, Sexton et al. (1991) and Baulch (1997) consider three distinct trade regimes: an "at the parity bounds" regime where the spatial price difference equals the unit intermarket arbitrage cost; an "inside the parity bounds" regime where the local prices differ by less than that cost; and an "outside the parity bounds" regime where the observed spatial price difference is larger than that arbitrage cost.

Our point of departure is the extension proposed by Negassa and Myers (2007) in their analysis of the impacts of a marketing policy change on agricultural markets integration in Ethiopia. By allowing possible dynamic shifts in regime probabilities, this enriched PBM framework makes it possible to assess whether the policy change modifies or not the probabilities of observing the various trade regimes. By construction, this model provides an adapted methodology to investigate our research questions. The present paper thus details the first application of this extended PBM to model an energy-related commodity. In addition, this paper offers an enriched specification as it shows how the extended PBM can be combined with the methodology in Kleit (2001) to correct for the impact of first-order autocorrelation on the efficiency of the estimates.

The remaining sections of this paper are organized as follows. Section 2 presents the background and the motivation of our analysis. Section 3 describes the econometric methodology. Then, section 4

[3] *To the best of the authors' knowledge, the literature on the integration of energy markets only provides three examples of empirical analyses based on the PBM methodology: Kleit (1998, 2001) and Bailey (1998).*

presents and discusses the empirical findings. Finally, the last section offers a summary and some concluding remarks.

2. Background and motivation

This section consists of three different parts. In the first subsection we briefly review how the US sanctions imposed on Iran's energy sector have shaped the emergence of an export-oriented petrochemical industry. A second subsection presents the last round of sanctions, discusses their consequences on the country's export behavior, and clarifies the motivation for the present analysis. The last subsection reviews the trade patterns of the international methanol market.

2.1 Iran's petrochemical strategy: Emergence of a methanol-exporting nation

In 1996 the US Congress enacted the Iran–Libya Sanctions Act (later renamed the Iran Sanctions Act) that gave the President the authority to impose sanctions on any company, organization or person (either domestic or foreign) that invested in Iran's petroleum sector. This act was aimed at reducing Iran's ability to fund the development of a nuclear program and had a profound impact on the country's strategy to monetize its vast hydrocarbon resources. By constraining foreign investment, it is reputed to have hampered the construction of new energy export infrastructures (e.g., pipelines or terminals for oil, natural gas pipelines, and LNG liquefaction plants).[4]

To partly circumvent these US restrictions, Teheran strongly encouraged the deployment of export-oriented resource-based industries, particularly in the petrochemical sector. The country's export of petrochemical products steadily increased from US $141 million in 2000 (i.e., less than 0.5% of the country's total annual exports) to $2.97 billion in 2010, representing 3.55% of the country's exports that year (U.N., 2015). Within this sector, the most significant achievement has been the rapid emergence of a world-class methanol industry that generated 34.8% of Iran's petrochemical export revenues in 2010.

The Iranian state capitalism indubitably triggered the emergence of the methanol sector as the state-owned National Petrochemical Company (NPC) played a decisive role in the design, construction, and operation of the country's major methanol complexes in both Bandar Imam and Assaluyeh (this 3.3 million-tons-a-year complex is the world's largest methanol plant). At least four lines of arguments explain the appeal of methanol processing for the Iranian state. First, methanol represents a profitable option to monetize natural gas resources. The mean-variance portfolio analysis in Massol and Banal-Estañol (2014) reveals that methanol provides the second-highest level of expected resource rents after LNG among the six main gas-based industries that can be implemented

[4] *As far as natural gas is concerned, a casual comparison with Qatar provides an instructive illustration of the impact of these restrictions. Both countries share the world's largest offshore gas field, the so-called South Pars/North Dome field. However, the extraction activities conducted in Iran have remained confined to rather preliminary stages of development compared to Qatar. This sharp difference is also noticeable when considering LNG projects. Qatar is now the world's largest LNG exporter whereas none of the four large LNG projects once envisioned in Iran (namely, Persian LNG, Iran LNG, Pars LNG and NIOC LNG) have been decided.*

within a gas-rich country. As the construction of an LNG liquefaction plant was impeded by the US sanctions, methanol represents an attractive second-best option.

Second, the global methanol market is experiencing steadily growing demand figures. During the last decade, world consumption grew at an average rate of +6% a year to attain 60.7 million metric tons by 2013. This raw material can be converted into a wide variety of products including formaldehyde (an important chemical raw material extensively used in particle board, plywood, paints, foams, rubbers, adhesive, coatings, resin plastic, explosives, pharmaceuticals, and pesticides), acetic acid and other petrochemicals (e.g., MTBE – a gasoline additive aimed at raising the octane number). In addition to these traditional chemical markets, the use of methanol as a fuel in the transportation sector is also rapidly emerging[5] and methanol is increasingly presented as a credible pathway for producing liquid fuels for spark-ignited engines.[6]

Third, compared to LNG, the methanol industry is less capital intensive and involves simpler processing technologies. These are well-suited features for the implementation of that industry in a country impacted by international sanctions.

Lastly, the exportation of gas-based methanol is less vulnerable to foreign sanctions than those of natural gas. The logistics are easier to handle as it can be transported using usual chemical tankers whereas LNG trade involves a tiny fleet of specific cryogenic vessels.[7] Moreover, as methanol's logistics involves far less specific assets, standard transaction cost economics suggest that the trade of methanol is less likely to be governed by complex long-term contracts that only a handful of purchasers are capable of signing. Accordingly, methanol exports are less likely to be controlled by a small set of foreign players, a feature that conceivably makes this trade less vulnerable to foreign sanctions.

Table 1. Iranian methanol and international trade

[PLEASE INSERT TABLE 1 HERE]

This industrial voluntarism has allowed Iran to rapidly gain control of a substantial share of the world trade of methanol (cf. Table 1). The volumes and trade value figures now position the country as the third largest exporter after Saudi Arabia and Trinidad and Tobago.

Given the scale attained by Iranian chemicals production and the specificity of the methanol industry, one may wonder whether the international sanctions decided in 2012 have impacted the Iranian methanol industry and transformed the worldwide methanol landscape.

[5] *In recent years, China has introduced two methanol-based fuels, the M100 (pure methanol) and M85 (the mass fraction of methanol is 85%) which are now distributed in Shanghai City, Shaanxi Province, and Shanxi Province (Su et al., 2013).*
[6] *As this commodity represents a credible alternative to oil in a number of uses, G. A Olah – a Nobel Laureate in Chemistry – is strongly advocating the emergence of a "Methanol Economy" (Olah, 2005).*
[7] *According to the enumerations detailed in BRS (2014) and GIIGNL (2014), only 393 ships are capable of providing LNG shipping services whereas the world chemical tankers tanker fleet includes about 2,460 vessels.*

2.2 Sanctions against Iran and their impact on petrochemicals exports

During the mid-2000s, a series of intelligence reports confirmed that Iran had resumed its nuclear weapons development program. Between 2006 and 2010, four sets of UN Security Council resolutions imposed progressively stronger sanctions focused on uranium and the nuclear sector. At the end of 2010 the unilateral and UN sanctions were reputed to remain far too weak to have a critical impact on Iran (Cordesman et al., 2014). Though their impact on the country's energy sector was limited, these multilateral decisions triggered the adoption of a new round of unilateral sanctions by both the US and Europe. The measures implemented between 2010 and 2012 focused on the most effective way to restrict trade with Iran and broadened the scope of the sanctions from a sectoral point of view. Then the United States decided to extend sanctions to all the economic and financial sectors in order to cut external funding sources and limit the development of Iranian energy companies. At the beginning of 2012, in order to isolate the Central Bank of Iran (CBI) and the country from international transactions, the US decided to exclude from the US banking sector any foreign banks which might have financial relations with the CBI. For its part, Europe also increased pressure on Iran with a new series of measures focused on two specific dimensions. The establishment of an oil embargo from the EU in January 2012[8] was accompanied by a technology sales ban for the oil, gas, and petrochemical sectors.

In March 2012 a set of further financial restrictions, including a prohibition for European companies to provide insurance tools for Iranian commercial transactions and more especially for transportation, was implemented. Furthermore, Europe isolated Iran from the international financial system by prohibiting access to the SWIFT interbank settlement system. During summer 2012 the US stepped up measures by expanding sanctions to all companies conducting business with Iran in all energy sectors, the insurance sector, and in transportation. Altogether, this batch of sanctions triggered a sharp increase in the insurance cost for Iranian tankers. In October 2012 the EU decided, among other decisions, to expand the ban on exports to Iran to other products that are key components for the energy sector: aluminum, graphite or steel and finally to all the equipment that can be used in the oil, gas, and petrochemical sector. Ultimately, these sanctions are reputed to have bitten hard, particularly by choking off access to technology, cheap shipping, and insurance. Indeed, for the latter around 90% of the world's tanker insurance is based in the West, so the sanctions threatened oil and petrochemical shipments to Iran's top Asian buyers China, India, Japan, and South Korea.

According to the US Energy Information Administration,[9] the Iranian crude oil exports decreased by 50% in 2012 compared to 2011 and Iran has tried, with mixed success, to mitigate the effects of sanctions. Iran developed a massive barter exchange system, especially with China, and is reputed to have bartered oil against Chinese consumer goods. According to industry sources, China also imported gas condensate from South Pars gas field in the same way. Then, up until 2012, market players developed very inventive means in order to maintain exchanges in petrochemicals and petroleum

[8] *In 2012, Iran exports to EU represent 23% of the global Iranian oil exports.*
[9] *EIA: "Iran's oil exports not expected to increase significantly despites recent negotiations" (December 10, 2013).*

markets as illustrated in ICIS's chemical analyses:[10] "*Some of the means cited by industry sources include shipping Iranian cargoes in vessels flying a different flag, or mixing the product with those of other countries, or hiding the origin of the cargo.*" Those ingenious means had consequences in terms of prices in the methanol market: Iran was forced to sell its own products below market prices so as to maintain export volume during the sanctions.

2.3 Methanol: Production, markets, and trade flow characteristics

In 2013 the global production of methanol reached around 61 million metric tons. China (43% of the world production), and more globally Asia (50%), is the largest producing region followed by the Middle East countries (20%), South America (12.5%), CIS and the Baltic States (6%), Africa (4%), Western Europe (4%), and North America (3%). Iran owns approximately 31% of the Middle East capacity for methanol (around 5 million metric tons) behind Saudi Arabia (45%).

International trade plays a crucial role in the methanol industry: the total exports of the commodity amounted to 25.5 million metric tons in 2013 (i.e., 42% of the world production that year). The Middle East region appears to be the first net exporting region with around 40% of the world exports for methanol and 10.2 million metric tons. South America (27% of global exports), Southeast Asia (12%), and Africa (9.5%) complete the picture of the exports (Figure 1). However, for Southeast Asia the situation appears to be more complex as the region is also a large importer with 8% of the total imports in volume. Some South Eastern Asian countries are net importers (e.g., Singapore, Thailand, Vietnam) whereas others are net exporters (e.g., Brunei, Malaysia). In term of consumption, China is the largest consumer (50% of the world consumption) and importer (18% of world imports). Asia as a region represents 46% of the world imports for methanol. The US (22%) or North America (24%) and Western Europe (18%) are also key players for this market (Figure 2).

Figure 1: Exports of methanol by regions in 2013 (in %)

[PLEASE INSERT FIGURE 1 HERE]

Figure 2: Imports of methanol by regions or countries in 2013 (in %)

[PLEASE INSERT FIGURE 2 HERE]

From an industrial organization perspective, the methanol industry is little concentrated and its industrial structure is reputed to be competitive. The world's largest player – Methanex Corp. – owns around 6% of the global capacity and the cumulated share of the 10 largest producers amounts to 27% (IHS Chemical, 2014).[11] Iran's methanol sector is dominated by the National Petrochemical Company which controls 4.5% of the world methanol processing capacity (i.e., 4.4 million metric tons per year).

[10] http://www.icis.com/resources/news/2012/08/30/9584929/tighter-sanctions-fail-to-deter-iran-methanol-exports-to-asia/

[11] *Historically, the world methanol industry was far more concentrated as Methanex used to be a dominant player prior to the early 2000s. By operating a series of large plants located all around the world (e.g., in Chile, Canada, Trinidad & Tobago, New Zealand), the company was even able to impose price discriminations depending on the destination (Asia, Europe, US). Nevertheless, Methanex Corp's ability to impose such spatial price discriminations faded away at the end of the*

Due to its sizeable production capacity and its geographic position, Iran is now considered to be a significant player. It should also be mentioned that, according to the industry's comments, Iran is considered to be a swing producer, selling its own product on a spot basis depending on the situations observed in the destination markets in Europe and Asia. Prior to the sanctions, this behavior was reputed to have helped to minimize regional price differentials and favored the emergence of an efficient international market. The question of the integration of the different regional methanol markets is crucial for our analysis. There is a very limited literature assessing the integration of methanol markets. To the best of our knowledge, there is just one study examining this topic: Mansur et al. (2010) examine the link and interaction between natural gas markets and the European, US, and Asian methanol markets. Using monthly Henry Hub and methanol prices during the period 1998–2007, they find that natural gas prices are co-integrated with methanol prices in the three regions. Furthermore, they estimate an error correction model and the findings provide evidence that natural gas prices drive methanol prices in Europe and the US, although this is not the case for Asia. This last finding is not so surprising as most of the methanol consumed in Asia is produced in plants where the methane feed gas is purchased through long-term contracts with specific (and mostly oil-indexed) pricing provisions.

Prior to the 2012 sanctions, the typical spatial organization of international methanol trade was as follows. The US, Europe, and China together represented around 63% of the world imports in 2011 (source: United Nations Commodity Trade Statistics Database). The methanol traded in the Atlantic basin is mainly exported from Trinidad and Tobago, Equatorial Guinea, and Venezuela. These three countries have the ability to realize some spatial arbitrage between the US and Western Europe as together they account for 91% of the US imports in volume and 18% of the EU imports in 2011. Europe is also supplied by methanol coming from Russia (15% of the EU imports), Saudi-Arabia (10%), Norway (7.5%), Iran (7%), and Northern Africa (Egypt, Algeria, Libya together represented 9%). The Middle East countries trade their methanol within the three main importing areas (the US, Europe, and Asia). For Middle East countries, Asia is a very interesting region for short-term arbitrage as China (20% of world imports), Japan (6.1%), South Korea (6%), Taiwan (4%) or India (4%) are eager to import methanol. Overall, the Southeast Asia region is a net exporter but the concurrency of a large export flow (around 3.0 million metric tons) from countries like Malaysia or Brunei and significant import flow (1.9 million metric tons) to countries like Singapore, Thailand or Vietnam indicates the possibility that arbitrage could be performed within the region.

3. Methodology

This section presents the empirical methodology used in this manuscript. We first review the specification of an adapted parity bounds model. Subsequently, we detail an extension to that model aimed at correcting for the possible presence of first-order serial correlation.

1990s – early 2000s with the emergence of new producers with large and modern plants (e.g., in Brunei, Equatorial Guinea, Iran, Qatar, Saudi Arabia).

3.1 An adapted parity bounds model

To begin with, we follow Negassa and Myers (2007) and detail an extension to the standard PBM proposed by Sexton, Kling, and Carman (1991).

We consider two markets i and j located in different regions that trade a homogeneous commodity. We aim to analyze the arbitrages that can be performed from market j to market i at time t and respectively denote P_{it} and P_{jt} the local market clearing prices in each location. The price variables are in levels and not in logarithms. We assume that there are no transport lags so that spatial arbitrage can take place within each observation period. The marginal arbitrage cost,[12] including all the transportation costs and the transaction costs incurred when performing an arbitrage from market j to market i at time t is denoted T_t .

The observed spatial price differentials and arbitrage costs can be used to define a taxonomy of three mutually exclusive trade regimes governing the arbitrages from market j to market i .

In regime I, the spatial price spread is equal to the marginal arbitrage cost. Thus, the following condition is binding:

$$P_{it} - P_{jt} - T_t = 0 . \tag{1}$$

As highlighted in Barrett and Li (2002), this regime is consistent with the conditions for perfect spatial market integration irrespective of whether trade occurs. When trade occurs the local market prices will differ from autarky prices and the supply and demand shocks observed in one market will be transmitted to the other market.

In regime II, the spatial price difference is greater than the marginal arbitrage cost:

$$P_{it} - P_{jt} - T_t > 0 . \tag{2}$$

In this regime, markets are separated and there are unseized opportunities for profitable spatial arbitrage. This regime can be generated by a host of factors including: the exertion of market power by the arbitragers, the implementation of governmental restrictions to trade, and the existence of capacity constraints on the transportation infrastructure.

In regime III, the marginal profit to arbitrage from j to i is strictly negative:

$$P_{it} - P_{jt} - T_t < 0 . \tag{3}$$

In this regime, there is no incentive to trade. If trade is not occurring, the observed local prices correspond to autarky prices. If trade is occurring, it provides negative profits, which is not consistent with spatial equilibrium conditions.

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[12] In this paper, we implicitly posit that the total arbitrage cost is a linear function of the quantities traded.

We now detail the empirical specification aimed at estimating the probabilities of being in each regime using a data set of N observations for the local market-clearing prices and the marginal arbitrage cost.

From an empirical perspective, the marginal arbitrage cost T_t is usually seldom available to the modeler. Hereafter, this arbitrage cost is assumed to be explained by a constant and by a vector of observable exogenous factors Z_t :

$$T_t = \alpha + Z_t \beta + \varepsilon_t \ , \tag{4}$$

where α and β are unknown parameters to be estimated, and ε_t is a random error that is assumed to be i.i.d. normally distributed with a zero mean and standard deviation σ_ε .

Using that specification, the conditions for the three regimes can be written:

Regime I: $$P_{it} - P_{jt} - \alpha - Z_t \beta = \varepsilon_t \ , \tag{5}$$

Regime II: $$P_{it} - P_{jt} - \alpha - Z_t \beta = \varepsilon_t + \mu_t \ , \tag{6}$$

Regime III: $$P_{it} - P_{jt} - \alpha - Z_t \beta = \varepsilon_t - \mu_t \ , \tag{7}$$

where μ_t is a random error that is assumed to be i.i.d. from a zero-centered normal distribution truncated above at 0 with a standard deviation parameter σ_μ. These three conditions are embedded within a switching regression framework that is estimated using a maximum likelihood method.

In a PBM, the goal is to determine the probabilities of being in each regime. In this paper, these probabilities are allowed to change following the sanctions. Thus, it is assumed that two periods can be distinguished within the sample depending on whether international sanctions were imposed on Iranian methanol or not. In each period, trading regime probabilities are constant over time but the probabilities of the two periods can differ.[13] Formally, we let λ_r denote the estimated probability to observe regime r before the policy change, δ_r measure the change in the probability of being in regime r due to the policy change, and D_t denote a dummy variable that takes a value of 1 after this change. Hence, the probability of being in regime r at time t is $\lambda_r + \delta_r D_t$. For notational simplicity, we let λ and δ denote the vectors of these parameters to be estimated.

[13] In this paper, we thus follow the approach of Park et al. (2002) and assume that the policy change induces a discrete and instantaneous jump in the probabilities of being in each regime. Technically, Negassa and Myers (2007) have proposed relaxing this assumption by modeling D_t as a transition variable and allowing an intermediary adjustment period during which its value linearly increases from 0 to 1. In our empirical application to the methanol markets, this refinement has been tested but was finally abandoned because of sample size considerations.

Denoting $\pi_t = P_{it} - P_{jt} - \alpha - Z_t \beta$ the random variable that gives the expected value of the difference between the spatial price spread and the marginal arbitrage cost at time t, the joint density function for the observation at time t is the mixture distribution:

$$\begin{aligned} f_t\left(\pi_t \middle| (\lambda,\delta,\theta)\right) \; = \; & (\lambda_I + \delta_I D_t)\, f_t^I\left(\pi_t \middle| \theta\right) \\ & + (\lambda_{II} + \delta_{II} D_t)\, f_t^{II}\left(\pi_t \middle| \theta\right) \\ & + (\lambda_{III} + \delta_{III} D_t)\, f_t^{III}\left(\pi_t \middle| \theta\right) \end{aligned} \tag{8}$$

where: $\theta = (\alpha, \beta, \sigma_\varepsilon, \sigma_\mu)$ is a parameter vector to be estimated, $f_t^I\left(\pi_t \middle| \theta\right)$ is the normal density function, and $f_t^{II}\left(\pi_t \middle| \theta\right)$ (respectively $f_t^{III}\left(\pi_t \middle| \theta\right)$) is the density function derived in Weinstein (1964) for the sum of a normal random variable and a centered-normal random variable truncated above (respectively below) at 0.

The likelihood function for a sample of observations $\{P_{it}, P_{jt}, Z_t\}$ is:

$$L(\lambda,\delta,\theta) = \prod_{t=1}^{N} f_t\left(\pi_t \middle| (\lambda,\delta,\theta)\right). \tag{9}$$

The model can be estimated by maximizing the logarithm of the likelihood function with respect to regime probabilities during the first period, the changes in the probabilities caused by the policy change, and model parameters subject to the following constraints: (i) the regime probabilities lie in the unit interval in each period (i.e.: $0 \le \lambda_r \le 1$ and $0 \le \lambda_r + \delta_r D_t \le 1$ for any regime), (ii) the regime probabilities sum to one in each period (i.e.: the restrictions $\sum_r \lambda_r = 1$ and $\sum_r \delta_r = 0$), and (iii) the standard deviations are non-negative.

This framework allows us to test the null hypothesis of no structural change in any regime probability. This can be done using a likelihood-ratio test approach based on the unrestricted model above and a restricted one with no structural change in the regime probabilities (i.e.: the probability to observe each regime is kept constant over time).

3.2 Correcting for autocorrelation

To the authors' knowledge, most previous parity bound models do not account for autocorrelation. Kleit (2001) is one of the few exceptions. Surprisingly, this omission is seldom discussed. Yet serial correlation due to both supply shocks and speculative storage activity is commonly observed in the empirical studies dedicated to commodity prices (Deaton and Laroque, 1996).[14] As the presence of unmodeled autocorrelation can result in inefficient estimates, the presence of serial correlation has to be appropriately corrected for.[15]

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[14] *In our empirical application to the methanol markets, the presence of autocorrelation can be justified by that industry's specific timeline. A chemical vessel is a slow-moving transportation system that typically needs at least a week to move a*

In this paper, we apply the Bayesian approach in Kleit (2001). We aim to extend the model above to adjust for the presence of serial correlation in the error term ε_t . Yet, a difficulty emerges: the exact value ε_{t-1} cannot be directly observed. However, we can consider the expected value of ε_{t-1}, given the evidence provided by the previous observation, which results in the modified specification:

Regime I: $$P_{it} - P_{jt} - \alpha - Z_t \beta - \rho \mathrm{E}\left(\varepsilon_{t-1} \middle| \eta_{t-1}\right) = \varepsilon_t , \tag{10}$$

Regime II: $$P_{it} - P_{jt} - \alpha - Z_t \beta - \rho \mathrm{E}\left(\varepsilon_{t-1} \middle| \eta_{t-1}\right) = \varepsilon_t + \mu_t , \tag{11}$$

Regime III: $$P_{it} - P_{jt} - \alpha - Z_t \beta - \rho \mathrm{E}\left(\varepsilon_{t-1} \middle| \eta_{t-1}\right) = \varepsilon_t - \mu_t . \tag{12}$$

where: ρ is an autocorrelation coefficient such that $-1 < \rho < 1$; η_{t-1} is the observed lagged residual, that is, $\eta_{t-1} = \pi_{t-1} - \rho \mathrm{E}\left(\varepsilon_{t-2} \middle| \eta_{t-2}\right)$; and $\mathrm{E}\left(\varepsilon_{t-1} \middle| \eta_{t-1}\right)$ represents the expected value of ε_{t-1} given evidence provided by the observed lagged residual.

We now explain how the expected value $\mathrm{E}\left(\varepsilon_{t-1} \middle| \eta_{t-1}\right)$ can be computed. Given the observed value of the lagged residual η_{t-1} and the parameter vector $\theta_1 = (\theta, \rho)$, Bayes theorem can be invoked to evaluate $\mathrm{P}_{t-1}^r = \mathrm{P}_{t-1}\left(r \middle| \eta_{t-1}, \theta_1\right)$ the posterior probability that the residual observed at time $t-1$ was generated by regime r (Kiefer, 1980; Spiller and Wood, 1988):

$$\mathrm{P}_{t-1}^r = \frac{\left(\lambda_r + \delta_r D_{t-1}\right) f_{t-1}^r\left(\eta_{t-1} \middle| \theta_1\right)}{\sum_{k \in \{I, II, III\}} \left(\lambda_k + \delta_k D_{t-1}\right) f_{t-1}^k\left(\eta_{t-1} \middle| \theta_1\right)} . \tag{13}$$

The expected value $\mathrm{E}\left(\varepsilon_{t-1} \middle| \eta_{t-1}\right)$ can be constructed from the observed residual η_{t-1} by removing the expected contributions of regimes II and III. We proceed as follows. First, we can remark that the observed residuals are equal to the expected value of the observed residuals given the observed residuals η_{t-1}=$\mathrm{E}\left(\eta_{t-1} \middle| \eta_{t-1}\right)$. Hence, the observed residuals equal the sum of the expected values of the residuals observed in each regime, given the observed residuals weighted by the posterior probability to observe each regime. Then, this equation can be rearranged to obtain $\mathrm{E}\left(\varepsilon_{t-1} \middle| \eta_{t-1}\right)$ as the observed residual η_{t-1} minus $\mathrm{E}(\mu)$ the expected value of the non-negative random variable[16] that is added to ε_{t-1} in regime II times the posterior probability to observe that regime, plus the expected value of the

given cargo of methanol from one of the markets under scrutiny to another. As a result, the observed monthly prices are likely to jointly represent the outcome of decisions taken both during the current month and the preceding one.

[15] *Barrett and Li (2002, footnote 3) mentioned the serial correlation issue and claimed that the Cochrane-Orcutt method could be used to correct for serial correlation. However, the distribution of the observed residuals is dramatically modified from one observation to the next in case of a regime switch. Therefore, one may question the validity of a Cochrane-Orcutt approach.*

[16] *Denoting ϕ the density function of the standard normal distribution and Φ its cumulative distribution function, this expected values is:* $\mathrm{E}(\mu) = \sigma_\mu \phi(0) / \left(1 - \Phi(0)\right)$.

random variable that is subtracted from ε_{t-1} in regime III times the probability to observe that regime, i.e.:

$$\mathrm{E}\left(\varepsilon_{t-1}\middle|\eta_{t-1}\right)=\eta_{t-1}-\mathrm{P}_{t-1}^{II}\mathrm{E}(\mu)+\mathrm{P}_{t-1}^{III}\mathrm{E}(\mu)\,. \tag{14}$$

The construction of $\mathrm{E}\left(\varepsilon_{t-1}\middle|\eta_{t-1}\right)$ can be nested within the likelihood specification above. So, the estimation proceeds again from a maximization of the log-likelihood function with respect to the regime probabilities λ and the parameters θ_1 subject to the preceding constraints and to the restriction $-1<\rho<1$.

4. Application

The model was applied to the major methanol importing markets to estimate the effects of the trade policy sanctions on spatial market efficiency. We first present the data before analyzing the estimates.

4.1 Data

We assembled time series data containing 60 monthly observations on prices and shipping costs. We consider the period covering January 2009 to December 2013. This starting date is imposed by the unavailability of methanol price data in India prior to 2009. The end of the period corresponds to the partial lifting of the sanctions. We further distinguish two subperiods: from January 2009 to March 2012 there were no sanctions against Iran's petrochemical exports whereas after April 2012 sanctions were implemented.

We use monthly transaction price data for methanol delivered in China, India, the EU (Rotterdam), US and South-Eastern Asian economies (Indonesia, Malaysia, Philippines, Singapore, Thailand, Vietnam) as published by Platts, a price-reporting service.[17] These prices are denominated in US dollars per metric ton. It should be noted that there are no quality differences among these markets as the traded product is supposed to verify the quality specifications adopted by the International Methanol Producers and Consumers Association. Following the discussion on trade characteristics in section 2.3, our discussion will be centered on the following market pairs: EU–India, China–India, South-East Asia–China, South-East Asia–India, US–EU, EU–China, and EU–South-East Asia.

To estimate the marginal arbitrage costs of shipping from one market to another, we use an explanatory variable Z_t : the time charter rates as published by a shipbroker following a monthly assessment of the daily charter price for a representative chemical carrier with a cargo capacity of 16,500 deadweight tons (dwt) denominated in thousand US dollars (BRS, 2014). This series is reputed

[17] *For South-Eastern Asian economies, the Platts database does not provide country-specific prices but simply provides a regional price assessment.*

to provide insights into the average spot freight rates incurred by chemical shippers all around the world.

Figure 3 provides plots of the price series. A visual inspection of these plots suggests that the trajectory of EU and US prices could differ from those observed in Asian markets when the sanctions on Iran are implemented. To corroborate this preliminary observation, Table 2 details the correlation coefficients between the local price series in both levels and first differences for the seven market pairs under scrutiny. Following Stigler and Sherwin (1985), the high values observed before the implementation of the sanctions could be interpreted as positive signs of global market integration. In contrast, the "degree" of that global integration seems to be weaker when the sanctions are imposed as we observe lower correlation values between Europe and the other markets. Nevertheless, there are inferential dangers in using such simple correlation measures to test for market integration. Hence, these signs of a possibly changing degree of market integration call for further examinations.

Figure 3. Data plots of the price series (USD/ton)

[PLEASE INSERT FIGURE 3 HERE]

Table 2. Correlation coefficients for the price series

[PLEASE INSERT TABLE 2 HERE]

Table 3 summarizes the descriptive statistics for the spatial price spreads and the time charter rates. Except for the "US–EU" case, the distributional properties of the spatial price spreads show moderate signs of non-normality. In contrast, the estimated first-order autocorrelation coefficients reveal clear evidence of positive serial correlation for all series. This finding is in favor of a dynamic specification able to correct for serial correlation. Unit root issues were addressed using the Lanne-Lütkepohl-Saikkonen (LLS) test that incorporates the possibility of a level shift in April 2012 when the sanctions were implemented (see: Saikkonen and Lütkepohl, 2002; Lanne et al., 2002). From the LLS test statistics, the null hypothesis that the series contains a unit root is systematically rejected at the 5% significant level for all series. Therefore, we proceed under the premise that all examined series are stationary.

Table 3. Descriptive statistics for the spatial price differences and the time charter rates

[PLEASE INSERT TABLE 3 HERE]

4.2 Estimation and empirical results

a – Estimation procedure

The estimation procedure involves the constrained maximization of a non-trivial log-likelihood function. This is a non-linear, non-convex, constrained optimization problem that has to be solved numerically using hill-climbing procedures.[18]

To obtain a feasible starting point, we first consider the simplest possible specification (i.e., omitting the exogenous explanatory variable Z_t and assuming zero values for the autocorrelation parameter). The converged solution for this restricted specification is then used as a feasible starting point for the unrestricted model.

b – Empirical results

Estimation results are reported in Table 4. This table details the estimates obtained for: the unobserved marginal transaction costs (α, $\beta_{Shipping}$), the autocorrelation parameter (ρ), the regime probabilities (λ's), the changes in the regime probabilities following the sanctions (δ's), the standard deviation parameters for the distributions ($\sigma_\varepsilon, \sigma_\mu$), and the chi-square statistics for the likelihood ratio test of the joint hypothesis of no change in regime probabilities (χ^2).

To begin with, we examine the autocorrelation and arbitrage cost parameters. The estimated autocorrelation coefficients ρ are significant at the 1% level and their values are positive with an order of magnitude similar to those of the series $\left(P_{it} - P_{jt}\right)$. This finding confirms the need to explicitly model serial correlation in our specification. Regarding arbitrage costs, we have dropped the variable Z_t for the pairs EU–US, EU–China and EU–South-East Asia as the estimates of $\beta_{Shipping}$ were not statistically significant at the 10% level. In three other cases we observe significant but negative values for $\beta_{Shipping}$ which is surprising because the estimated marginal arbitrage costs are positive. Overall, these findings suggest that the time charter rates represent a mere proxy for the true unobservable transfer costs incurred by the arbitragers. Indeed, this exogenous variable suffers from a series of limitations. First, this series may be cursed by data quality issues as it is derived from an opaque monthly assessment conducted by a shipbroker that does not document its assessment methodology. Second, one needs to keep in mind that the size of an ocean-going tanker can range between 5,000 to 50,000 dwt whereas the assessment is conducted for a 16,500 dwt chemical tanker. Because of the heterogeneity of the world fleet, substantial variations in the effective unit charter rates incurred by arbitragers can exist. Lastly, this index has a global nature and thus cannot exactly account for the local factors modifying the tanker supply and demand conditions along each trade route (e.g., tanker availability, weather-related issues).

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[18] All the estimates reported in this paper have been obtained using STATA and an iterative procedure that performs 20 iterations using the Davidon-Fletcher-Powell (DFP) algorithm followed by 20 iterations using the Broyden-Fletcher-Goldfarb-Shanno (BFGS) one, followed by two iterations using the Newton-Raphson methods and then a switch back to DFP for 20 iterations, and so forth.

With these caveats in mind, we proceed and examine the regime probabilities. These results convey rich information on the extent of market integration both before and after the sanctions. Before the sanctions, there are very high estimates for λ_I in six of the seven market pairs. These estimates generally corroborate the observations derived from Stigler and Sherwin's (1985) correlation measures as they reveal that a high degree of integration was achieved among the three Asian markets and between each of them and Europe. In contrast, the EU and US markets for methanol are less perfectly integrated as the probability to observe a transatlantic spatial price spread lower than the marginal transfer cost is 0.31 between Europe and the US.

After the sanctions, we observe a complete reconfiguration of the spatial extent of these methanol markets. To begin with, it seems that the sanctions increase the degree of market integration across the Atlantic. The post-sanction probability to observe the "autarky" regime is zero whereas there is a net increase in the probability to observe the "at the parity bounds" regime I. One may object to the limited statistical significance of δ_I in that interpretation. Nevertheless, the probability shift is confirmed by the LR test since the null hypothesis "zero probability changes after the sanctions" is rejected at the usual 5% level.

Table 4. Estimation results for spatial arbitrages

[PLEASE INSERT TABLE 4 HERE]

Then, we notice that markets have also become highly segmented between Europe and each of the three Asian markets following the sanctions. It seems that there is a relative price increase in Europe (compared to the Asian markets) as the results reveal massive and highly statistically significant shifts in probabilities from regime I to regime II, i.e., from the "at the" to the "outside the parity bounds" regime that indicates the presence of barriers to trade. For the market pairs EU–India and EU–China, this shift has a radical nature as the likelihood ratio (LR) tests firmly reject the absence of probability changes after the sanctions. Interestingly, this hypothesis is only mildly rejected for Europe and South-East Asia, which suggests that the latter market may no longer be fully integrated with those in India and China. To clarify this, it is instructive to further examine the geographical extent of market integration within Asia following the sanctions. Between China and India, the LR test fails to reject the validity of the restricted model with unchanged probabilities at the 10% level which reveals that the sanctions have not modified the high degree of market integration observed between the two importing countries. An explanation of this finding is given by the similar nature of the policy responses implemented in both countries: India and China are reputed to have offered tankers sovereign guarantees to importers in order to maintain the importation of Iranian methanol after the sanctions. In contrast, it seems that by restricting access to shipping insurance services, the European sanctions have had a different price impact in South East Asia (i.e., Indonesia, Malaysia, Philippines, Singapore, Thailand, Vietnam) where governments have not offered such sovereign guarantees. Indeed, for the market pairs associating South East Asia with either China or India, there are total shifts in the regime probabilities from the "integrated" to the "barrier to trade" regime and the high values of the LR test statistics indubitably confirm the occurrence of a structural change.

5. Concluding remarks

This paper sheds light on the effects of the international sanctions against Iran and more particularly on their impacts on spatial integration of commodity markets. Using a case study focusing on the international methanol markets during the period 2009–2013, this paper shows that the extended parity bounds model, as presented in Negassa and Myers (2007), can be applied to derive a series of original findings.

Prior to the EU sanctions, a high degree of market integration was achieved between non-US markets. In contrast, we observe a complete reconfiguration of the spatial extent of these methanol markets under the sanctions as the markets became more fragmented. These findings corroborate the casual commentaries of industry observers who argue that the sanctions only imperfectly prevented the exportation of Iranian methanol to China and India as these two countries were reputed to have offered alternative insurance and transportation schemes that somehow alleviated the sanctions. Overall, these results document the importance of Iran as a swing producer integrating the global methanol markets.

As an extension for future research, it would be interesting to include monthly trade volumes – which are scarcely available – in addition to monthly price data to fully capture the phenomenon.

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Tables and Figures

Table 1. Iranian methanol and international trade

Year	Importations of Iranian methanol: Volumes (1,000 metric tons)	Share of globally imported volumes of methanol (%)	Trade Value (million USD)	Share of methanol's world trade value (%)
1995	29.8	0.3	11.2	0.4
2000	152.6	0.9	25.4	1.0
2005	1 021.6	5.0	267.7	4.8
2008	2 525.5	10.9	1 105.4	11.1
2010	3 975.9	14.4	1 171.0	14.8

Source: United Nations Commodity Trade Statistics Database (U.N. COMTRADE import database).

Table 2. Correlation coefficients for the price series

	EU–India	China–India	S.E. Asia–China	S.E. Asia–India	US–EU	EU–China	EU–S.E. Asia
In levels							
Entire sample	0.855*	0.976*	0.982*	0.967*	0.953*	0.891*	0.918*
Subperiod 1	0.910*	0.986*	0.986*	0.986*	0.917*	0.904*	0.903*
Subperiod 2	0.450*	0.897*	0.936*	0.866*	0.927*	0.637*	0.774*
In first differences							
Entire sample	0.291*	0.826*	0.827*	0.781*	0.134	0.382*	0.307*
Subperiod 1	0.910*	0.986*	0.986*	0.986*	0.917*	0.904*	0.903*
Subperiod 2	0.130	0.731*	0.798*	0.700*	0.076	0.369	0.328

Note: The variables are in levels and not in logarithms. Subperiod 1 is defined as the period covering January 2009 to March 2012. Subperiod 2 is defined as the period covering April 2012 to December 2013 (i.e., when the sanctions are implemented). * Significance at the 5% level.

Table 3. Descriptive statistics for the spatial price differences and the time charter rates

	$\left(P_{it} - P_{jt}\right)$							Z_t
	EU–India	China–India	S.E. Asia–China	S.E. Asia–India	US–EU	EU–China	EU–S.E. Asia	Shipping rates
Mean	50.336	30.886	4.232	35.118	52.040	19.450	15.218	9.534
Standard Deviation	47.997	15.447	23.574	26.045	32.034	43.397	36.398	1.768
Skewness	0.531	-0.384	0.577	0.765	-1.293	.444	0.179	1.840
Kurtosis	2.178	3.652	3.635	2.939	5.176	2.491	2.214	5.371
Jarque-Bera	4.503 *(0.105)*	2.538 *(0.281)*	4.342 *(0.114)*	5.867 *(0.053)*	28.553* *(0.000)*	2.624 *(0.269)*	1.863 *(0.394)*	47.912* *(0.000)*
Observations	60	60	60	60	60	60	60	60
First-order autocorrelation	0.853*	0.625*	0.802*	0.775*	0.695*	0.865*	0.757*	0.911*
LLS test	-2.960* (2)	-3.215* (1)	-3.281* (2)	-3.429* (0)	-2.897* (1)	-3.318* (2)	-3.510* (1)	-3.988* (1)

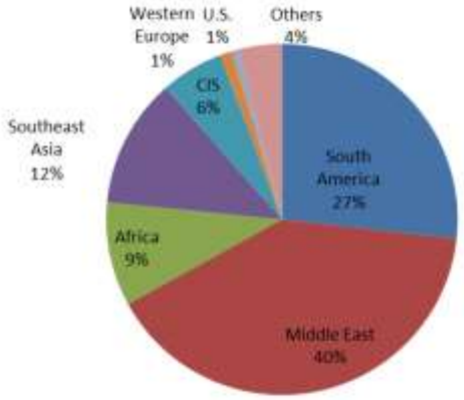
Note: The series are in levels and not in logarithms. * indicates significance at the 0.05 level. Numbers in parentheses are *p*-values for the null hypothesis of normality for the Jarque-Bera test. The null hypothesis for the Lanne-Lütkepohl-Saikkonen (LLS) test is "the series has a unit root I(1)." For the LLS test, the break date is known ("April 2012") and lag lengths are in parentheses. Critical values for the LLS test are tabulated in Lanne et al. (2002).

Table 4. Estimation results for spatial arbitrages

	EU–India	China–India	S.E. Asia–China	S.E. Asia–India	US–EU	EU–China	EU–S.E. Asia
Mean parameters							
α	71.885**	7.149	38.222**	51.965**	60.300**	2.333	6.673
	(26.299)	*(11.473)*	*(11.746)*	*(12.692)*	*(5.471)*	*(6.276)*	*(5.414)*
$\beta_{Shipping}$	-4.244	2.218	-4.316**	-2.839*			
	(2.561)	*(1.150)*	*(1.209)*	*(1.287)*			
ρ	0.608**	0.558**	0.691**	0.537**	0.474**	0.698**	0.706**
	(0.117)	*(0.118)*	*(0.096)*	*(0.116)*	*(0.100)*	*(0.084)*	*(0.078)*
Standard deviations							
σ_{ε}	23.106**	10.681**	9.780**	11.887**	18.344**	21.285**	22.993**
	(3.786)	*(1.339)*	*(1.051)*	*(1.423)*	*(2.284)*	*(3.537)*	*(2.344)*
σ_{μ}	53.775**	10.846	23.550**	32.529**	42.114*	44.860**	27.025**
	(12.471)	*(6.222)*	*(5.464)*	*(9.436)*	*(17.356)*	*(10.888)*	*(9.782)*
Probabilities							
λ_I	0.976**	1.000	1.000	1.000	0.687**	0.969**	0.972**
	(0.049)				*(0.217)*	*(0.043)*	*(0.061)*
λ_{II}	0.024	–	–	–	–	0.031	0.028
	(0.049)					*(0.043)*	*(0.061)*
λ_{III}	–	–	–	–	0.313	–	0.000
					(0.217)		*(0.008)*
Policy effects							
δ_I	-0.976**	-0.990**	-1.000	-1.000	0.313	-0.969**	-0.972**
	(0.049)	*(0.017)*			*(0.217)*	*(0.043)*	*(0.061)*
δ_{II}	0.976**	0.883**	1.000	1.000	–	0.969**	0.972**
	(0.049)	*(0.204)*				*(0.043)*	*(0.061)*
δ_{III}	–	0.108	–	–	-0.313	–	-0.000
		(0.187)			*(0.217)*		*(0.008)*
Log likelihood							
Unrestricted	-286.224	-233.284	-232.863	-247.060	-274.692	-279.737	-277.951
Restricted	-295.372	-234.407	-244.411	-259.022	-278.535	-287.759	-280.459
LR test							
$\chi^2(2)$ statistics	18.295	2.246	23.096	23.923	7.686	16.042	5.017
	(0.000)	*(0.325)*	*(0.000)*	*(0.000)*	*(0.021)*	*(0.000)*	*(0.081)*
Observations	59	59	59	59	59	59	59

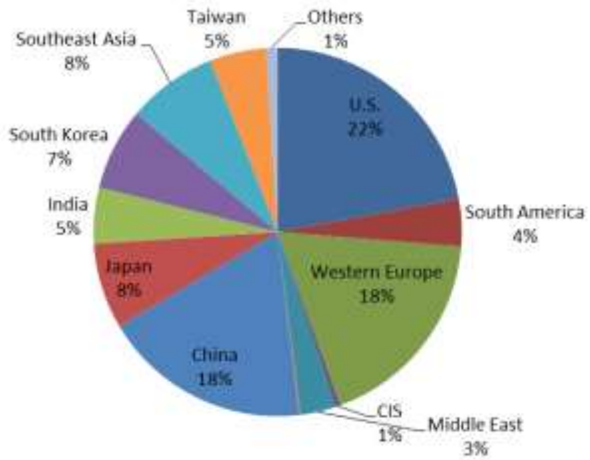
Note: Asterisks indicate significance at 0.05* and 0.01** levels, respectively. Numbers in parentheses are standard errors (*p*-values in the case of χ^2 statistics). A dash for the standard error indicates not calculated due to the parameter estimate being at the boundary of the parameter space. The models have been estimated using a general-to-specific approach by first including shipping rates as an explanatory variable and then dropping it and re-estimating if that variable was not significant at the 10% level.

Figure 1: Exports of methanol by regions in 2013 (in%)



Source: IHS Chemical (2014)

Figure 2: Imports of methanol by regions or countries in 2013 (in%)



Source: IHS Chemical (2014)

Figure 3. Data plots of the price series (USD/ton)