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Trade sanctions and international market integration: Evidence from the sanctions on Iranian methanol exports

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Abstract

This paper examines the impact of trade sanctions, imposed against large exporting nations, on the degree of spatial integration achieved between non-sanctioned importing markets. The analysis is conducted under a parity bounds framework based on Negassa and Myers (*American Journal of Agricultural Economics*, 89, 2007, 338). We apply this model to investigate the effects of the 2012–2016 sanctions against Iran's petrochemical exports on the main importing markets in Asia and we use it to measure the degrees of spatial integration attained outside and during the sanction period. Our findings document a complete reconfiguration of the spatial extent of the methanol markets. Outside of the sanction period, a high degree of market integration was achieved among the main Asian markets. In contrast, we observe the emergence of two little integrated market areas, China and India on one side and South Korea and South-East Asia on the other, when sanctions are imposed.

KEYWORDS

Iran, law of one price, market integration, methanol, sanctions

1 | INTRODUCTION

The use of economic sanctions, by nations or international organisations, aimed at influencing the decisions of another country, is now a common and recurring feature in international relations. Notorious examples include the sanctions imposed by the West on Russia in response to the invasion of Ukraine and the long series of nuclear-related sanctions targeting Iran's exports that are discussed in this paper. Independently of the political goal pursued (e.g. territorial disputes, nuclear non-proliferation, democracy, human rights, freeing captured citizens, environmental motives), the economic sanctions often aim at restricting the coerced country's access to international trade (Hufbauer et al., 2009; Lance & Engerman, 2003). Sanctions may include restrictions on the goods and services imported from, or exported to, the target country (e.g. a total blockade on the trade of some or all commodities or a set of selective restrictions on the provision of specific equipment or technologies).

An important strand of the economics literature has examined the impact of these trade sanctions on the sanctioning economy and/or the coerced one.[1] But, surprisingly, the consequences for third-party countries have so far attracted much less attention.[2] Still, sanctioning the exports of a large exporter is likely to provoke an important reorganisation of the international trade flows, even between third-party countries that are neither sanctioning nor being sanctioned. For instance, if the target country is a large exporter of a particular commodity, the economic sanctions imposed by a nation or a set of nations may conceivably affect the trading of this commodity in other importing markets.[3]

The purpose of this paper is to examine the impact that export restrictions may have on the degree of spatial integration of international markets. The law of one price states that, in integrated markets, homogeneous goods sold in different locations must sell for the same price (except for transportation costs). For example, in the case of energy and petrochemical commodities, the nearly continuous flow of sea-based trade from exporters to importers connects destination markets and links their pricing. The efficient arbitrage response to a *ceteris paribus* price increase in an importing market ‘A’ can involve rerouting some of the oceangoing tankers initially directed to a neighbouring importing country ‘B’ to ‘A’. Such redirections obviously

[1]The majority of this research has an empirical nature. Many contributions have analysed how these episodes of sanctions have adverse outcomes on the coerced nation, such as, for example, public health (Ali & Shah, 2000; Gibbons and Garfield, 1999; Aloosh et al., 2019), GDP growth (Yang et al., 2009), national currency (Peksen & Son, 2015) or income distribution (Afesorgbor & Mahadevan, 2016). Other empirical analyses explore whether threatened sanctions differ from imposed sanctions, compare the different instruments employed and whether their effect is product-specific (Afesorgbor, 2019). Another important strand of the literature surveys the historical use of economic sanctions to identify stylised features of their effectiveness (Hufbauer et al., 2009; Lance & Engerman, 2003). A few theoretical contributions have also emerged. Eaton and Engers (1992, 1999) propose a game theoretic analysis of the interactions between a sanctioning country and a target. Kaempfer and Lowenberg (1988) use a public choice perspective to model the adoption of sanctions within the sanctioning country (respectively, the reaction to the sanctions within the targeted country) as the outcome of the interactions among domestic interest groups competing to generate political pressure.

[2]To the best of our knowledge, only a handful of studies have investigated that issue. Notable exceptions are the empirical analyses of Caruso (2003) and Yang et al. (2009) who estimate a panel gravity model to examine the impacts of the sanctions on bilateral trade either between the sanctioning and the sanctioned nations or between the target and third countries.

[3]An example is given by the US decision to re-impose sanctions against Iran in 2018 that have triggered controversies in energy importing nations such as Japan (see *e.g.* Sakanashi et al., 2018).

affect the prices formed at both of these destination markets and are performed up to the point whereby a zero marginal profit is obtained by each arbitrager at each destination market. These spatial arbitrages are thus central to ensuring an efficient supply of the commodity. Yet, in the presence of trade sanctions, the arbitrage activity from the product of the sanctioned nation is restricted, potentially affecting market efficiency.

To explore whether, and how, the product allocation from producing to consuming countries is affected by trade sanctions, this paper proposes the use of an empirical approach based on the parity bounds model (PBM), first introduced in Sexton et al. (1991), that is widely applied in agricultural and energy economics to investigate either food security (e.g. Barrett & Li, 2002; Baulch, 1997; Zant, 2013) or energy issues (e.g. Massol & Banal-Estañol, 2018).[4] More specifically, we consider the extended version of the PBM proposed by Negassa and Myers (2007) that allows us to test for the effects of exogenous policy interventions on the degree of market integration. In a PBM, arbitrageurs are assumed to be profit-maximising agents. Using that assumption, intermarket price spreads are examined using a regime switching specification which estimates the probability of observing each of three possible 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 the arbitrage cost. By allowing possible dynamic shifts in regime probabilities, the extended PBM in Negassa and Myers (2007) makes it possible to assess whether the probabilities of observing the various trade regimes are affected by an exogenous policy change. By construction, this model provides an adequate methodology to investigate whether trade sanctions have affected market efficiency and, in particular, the observed spatial price spreads.

We consider as an application the unprecedented wave of sanctions imposed on Iran's hydrocarbon exports between 2012 and 2016 and examine their impact on price formation at destination markets. Iran is an appropriate case to conduct that investigation for at least two reasons. First, the country is a large exporter of these commodities to many destination countries in Asia. As the sanctions were predominantly imposed by Western nations, these sanctions could

[4]These PBM studies are part of the large empirical literature that, following Stigler and Sherwin's (1985) definition of an integrated market, investigates the spatial extent of a market for a homogeneous product traded in geographically separated markets. From a methodological perspective, these studies typically use local price data and apply either PBM or time-series techniques to measure whether the law of one price (LOOP) holds or not. Time-series analyses typically concentrate on the attention to the co-movement of prices for the study of market integration. Accordingly, a high degree of correlation and/or co-integration between the price series is interpreted as evidence that the law of one price is being enforced through spatial arbitrages. These empirical models provide useful insights into how local price shocks are transmitted to adjacent markets. However, this co-movement approach also has several limitations. First, the time-series techniques require trade to be conducted in a continuous manner with a direction that is fixed over time (Zant, 2013). Second, the time-series approach overlooks the role of transaction costs that can fluctuate independently from producer prices (Fackler & Goodwin, 2001). On that point, the Monte Carlo simulations conducted by Baulch (1997) and McNew and Fackler (1997) show that, because of the disregard for transaction costs, time-series models can generate flawed inferences. Lastly, time-series approaches have an 'atheoretical' nature as the modelling is chiefly guided by the desire to empirically capture the stochastic properties of the data-generating process at hand with no considerations for the underlying microeconomics (Zant, 2013). In the present study, these limitations can hardly be overlooked for the following reasons. The direction of trade is not constant as the professional literature recurrently presents Iran as an exporter that shifts the final destination of its petrochemical exports over time (IHS, 2017). The transaction costs incurred by methanol traders are likely to vary over time and sanctions can possibly have resulted in discontinuities in trade. In the present paper, we thus opt for the PBM approach that is theoretically consistent with the Enke–Samuelson–Takayama–Judge conditions for spatial price equilibrium (Zant, 2013).

possibly have had different impacts at the destination markets. Second, the detailed sanctioning measures imposed against Iran during the years 2012–2016 were largely unprecedented and largely inspired the sanctions imposed against Russia in 2022. These sanctions prohibited Iran's access to western-controlled services (e.g. marine insurance, banking system), to lines of credit for moving cargo and to fuel supplies for Iranian ships. Therefore, understanding the effects of these measures may provide useful insights into the future impacts of similar measures imposed against other sanctioned nations.

We focus on Iran's petrochemical sector, an industry which is the nation's second largest exporting sector after petroleum[5] and represented 2.7% of the country's export revenues in 2011 when the sanctions were conceived. Within that sector, we concentrate on Iran's largest petrochemical export: methanol — a basic petrochemical predominantly produced from natural gas, that is either used as a precursor to produce a variety of chemicals[6] or consumed as a fuel — because it is an homogeneous commodity that is traded in the main importing markets[7] and its local prices are denominated in a common currency. We focus on the destination markets in Asia because that region accounts for about 70% of the global consumption and Iran is reputed to play a special role in that regional trade. Indeed, in the professional literature, petrochemical analysts recurrently describe Iran as a large exporter which optimises its shipments between destination markets in Asia and thereby contributes to regional price integration (IHS, 2014, 2017).

The tradability of methanol and many other commodities can be impacted by a number of factors, such as periodic transportation bottlenecks. Hence, the spatial arbitrages connecting destination markets may show discontinuities. We first examine the effects of the sanctions on the observed spatial price gaps using standard linear specifications with time-invariant parameters. But, these preliminary investigations do not account for the possible presence of trade discontinuities (see the comments in Barrett, 1996, 2001; Baulch, 1997; McNew & Fackler, 1997). For that reason, we apply the PBM technique as the main analysis.

Our findings indicate that, outside of the sanction period, a high degree of market integration was achieved among the main Asian markets. In contrast, we observe a complete reconfiguration of the spatial extent of these methanol markets under the sanctions, as they became more fragmented. The disintegration forms two distinct market areas, respectively, comprising China and India on the one hand, and Korea and Southeast Asia on the other hand. The degree of market integration achieved within each of these two areas remains very high, as we find a high probability of these market pairs being 'at the parity bound'. In contrast, that probability is very low for the market pairs involving countries in each of the two market areas, as we find a high probability of outside the parity bound, which can be interpreted as objective signs of 'balkanisation.'

[5]In Iran, the petroleum sector accounts for 81% of the nation's export revenues. Yet, crude oil is poorly adapted to test for the impacts at destination markets because that trade is primarily channelled by long-term contracts with pricing clauses that are not disclosed but typically stipulate an indexation on publicly available price references (e.g. the price of Brent, WTI or Dubai grade). As a result, there are no markets for Iranian crude oil at destination countries. Hence, our decision to concentrate on petrochemicals can be described as a second-best option.

[6]The list of possible uses includes the conversion into formaldehyde (a raw material used in particle board, plywood, paints, foams, rubbers, adhesive, coatings, resin plastic, explosives, pharmaceuticals and pesticides), acetic acid, olefins (ethylene, propylene) or gasoline additives.

[7]These are marked differences from crude oil — another important Iranian export — which is not traded at destination and has a specific grade.

Under a perfectly functioning blockade of Iranian exports, all markets should have been symmetrically impacted by the privation of Iranian supplies. As a result, all the market pairs should have exhibited a lower degree of market integration than the one prevailing without the sanctions. But we observe a widening of the price differentials observed between the destination markets that are reputed to have imported Iranian products (China and India) and the rest (Korea, South-East Asia), whereas the price differentials within each of these two distinct market areas remained narrow, especially within the first. These findings are consistent with the opinions of market commentators who assert that Chinese and Indian importers obtained insurance from domestic providers to cover tankers carrying Iranian petrochemicals, whereas other non-sanctioned Asian countries stopped the importation of Iranian products and had to be supplied by producers located in other exporting countries (e.g. Saudi Arabia, Brunei).[8]

This paper contributes to the literature on the effects of the trade sanctions by investigating the impact of sanctions on prices formed in third-party countries. We provide evidence of the important reorganisation of the international trade flows following the sanctions. Sanctions may affect the degree of international market integration and market efficiency. We also show that the effects can be asymmetric, depending on each of the importing countries' abilities to circumvent the trade sanctions.

This paper also contributes to the small, and very much needed, literature attempting to shed a light on the consequences of the trade sanctions imposed on Iran. Regarding public health, Aloosh et al. (2019) study the multiple adverse consequences on population and individual health in Iran and analyse how the sanctions have affected health-care delivery, access to care, as well as their negative impacts on the social determinants of health. In economics, the existing theoretical literature gathers two analytical contributions: Naghavi and Pignataro (2015) formally examine the interactions between the sanctions and the theocratic regime's behaviour, and Miyagiwa and Ohno (2015) assess the sanctions' potential to stop Iran's nuclear programme. Gharibnavaz and Waschik (2018) use a calibrated computable general equilibrium (CGE) model to simulate the effects of international sanctions on the sanctioned economy and study their domestic redistributive effects. According to the simulation results, the Iranian economy incurred a reduction of the country's aggregate welfare of 14–15%, a loss that particularly affected the poorest urban and rural households and the Iranian government. Three other empirical contributions have also investigated the macroeconomic repercussions on Iran. Gharehgozli (2017) and Ghorbani Dastgerdi et al. (2018) investigate the impact on Iran's real GDP and inflation respectively. Dizaji and van Bergeijk (2013) model the interplay of macroeconomic and political variables and their dynamics.

To the best of our knowledge, the consequences on international trade have been barely studied so far. A notable exception is Haidar (2017) who first investigated the reaction of private Iranian exporters. His approach builds upon the ideas in Bown and Crowley (2007) who first defined 'export deflection' as a change in the destination of exports in response to an increase in a trade barrier in another market. Using a panel of private firms, his analysis shows that a majority of Iranian private exports were deflected to non-sanctioning countries. However, that study focuses on the private sector which de facto excludes the Iranian oil, gas and petrochemical

[8]Sources: https://www.insurancejournal.com/news/international/2012/05/17/247850.htm retrieved on March 17, 2023; https://www.reuters.com/article/iran-shipping-china/china-shippers-profit-from-sanctions-on-iran-petchem-trade-idUSL4E8G411O20120504 retrieved on March 17, 2023. The behaviour of China may be explained by the geopolitics and in particular by its strained diplomatic relations with the West.

sectors that are state controlled. So, the impacts on the trade energy-related commodities have so far not been analysed, which is quite astonishing given Iran's importance on the global energy scene and thus in the public policy debates of energy importing nations. One of the goals of this paper is to fill this gap.

The rest of the paper is structured as follows. The next section briefly describes the Iranian methanol industry, the organisation or the international trade of methanol, and the sanctions imposed on Iran so as to further clarify the background for our analysis. Section 3 presents the econometric methodology. Then, Section 4 presents the data and Section 5 discusses the empirical findings. Finally, the last section offers a summary and some concluding remarks.

2 | BACKGROUND

This section provides a brief review of Iran's export-oriented petrochemical strategy and how the sanctions imposed in 2012 have impacted the country's exports.[9] It also provides a condensed overview of the structure and organisation of the international trade of methanol.

2.1 | Iran and the global methanol trade

Since 2000, Iran has been developing a large-scale, export-oriented, petrochemical industry to monetise its vast natural gas resources. The country's export of petrochemical products steadily increased from US $141 million in 2000 to $2.97 billion in 2010 (UN, 2022). Within the petrochemical sector, the most significant achievement has been the rapid emergence of a world-class gas-based methanol processing industry, which is proving to be highly lucrative.[10] The state-controlled National Petrochemical Company (NPC) is now the world's second largest producer with an annual production capacity of 5 million tons. The NPC's exports are predominantly directed to Asia as these destination markets have a limited geological endowment in natural gas which imposes the importation of methanol from foreign sources (IHS, 2017).

The methanol trade is channelled through a series of short-term spot markets for methanol that are supply–demand driven. These markets are located in the most important importing markets and thus provide market clearing prices for standardised cargoes to be delivered to that specific destination. It is also important to highlight that, in contrast to most petroleum derivatives, methanol can be considered as a homogeneous good: It is a globally standardised product and there are no regional variations in quality standards.

Lastly, it is important to stress the NPC's behaviour in the global trade of methanol. The company is reputed to shift petrochemicals to destination markets in Asia that offer the highest netback price, thereby contributing to price integration among these markets absent any sanctions (IHS, 2017).

[9]For concision, the present section concentrates solely on the international sanctions imposed on Iran between 2012 and 2016. We refer to Torbat (2005) for an overview of the preceding sanctions episodes that were unilaterally imposed by the United States after 1979.

[10]Massol and Banal-Estañol (2014) note that the conversion of natural gas into methanol provides the second-highest level of expected resource rents after LNG among the main gas-based industries that can be installed in a gas-rich nation.

2.2 | Sanctions against Iran's petrochemical sector

The aspiration to set stringent limits on Iran's nuclear activities and the broader prospects for the accommodation of regime change have motivated a number of economic sanctions. In 2012, the advancement of Iran's nuclear activities convinced the global powers to increase global pressure via a new round of sanctions.[11] To isolate the country from international transactions, the United States decided to exclude from the US banking sector any foreign banks which might have financial relations with the Central Bank of Iran and the EU prohibited Iran's access to the SWIFT interbank settlement system. In addition, the EU placed an embargo on the importation of Iranian oil and a technology sales ban for the hydrocarbon and petrochemical sectors. In March 2012, the EU's measures were completed by a series of financial restrictions that included a prohibition for European insurers and reinsurers from covering tankers carrying Iranian oil and petrochemicals. This last measure came into effect on May 1. These sanctions were lifted in 2016 when Iran and global powers implemented the so-called Joint Comprehensive Plan of Action (JCPOA) that curtailed the country's nuclear weapons programme in exchange for sanctions relief.

Altogether, this batch of sanctions was aimed at choking off the petrochemical sector's access to technology, cheap shipping, insurance and banking. Indeed, while most of the Iranian petrochemical shipments were directed to Asia, the vast majority of the world's tanker fleet, including those operated by Iran's Petrochemical Transportation Company, is covered by Western-based protection and indemnity clubs, which insure against personal injury and environmental clean-up claims.

The Iranian reaction to these sanctions involved a reconfiguration of the country's export structure. The annual trade statistics reported in the UN Comtrade database indicate that, under the sanctions, Iran's exports of methanol to some countries such as South Korea vanished whereas the flows directed to China and India increased substantially. Market commentators argue that Chinese and Indian importers have obtained insurance from domestic providers and reaped huge profits from conducting that niche trade.[12] Yet, it remains unclear whether these official statistics accurately capture the effects of that export deflection as Iran is suspected to have backed the development of alternative and unconventional trading schemes. Iranian officials declared that boycotting Iran in the petrochemical market results in an encouragement of the black market.[13] Professional publications reported signs of an intensification of methanol smuggling activities that involved shipping Iranian cargoes in vessels flying a different flag, or mixing the product with those of other countries through middle-of-the-night, ship-to-ship transfers (Cordesman et al., 2014) or hiding the origin of the cargo.[14] The extent to which that reconfiguration and these unconventional trading schemes have affected destination markets and their degree of spatial integration remains to be examined.

[11]We refer to Patterson (2013) for a comprehensive chronological presentation of the EU sanctions against Iran.

[12]Sources: https://www.insurancejournal.com/news/international/2012/05/17/247850.htm retrieved on March 17, 2023; https://www.reuters.com/article/iran-shipping-china/china-shippers-profit-from-sanctions-on-iran-petchem-trade-idUSL4E8G411O20120504 retrieved on March 17, 2023.

[13]Source: the interview given on June 3, 2013 by the CEO of the National Iranian Petrochemical Company – http://www.farsnews.com/newstext.php?nn=13920312000249 retrieved on March 18, 2019.

[14]Source: http://www.icis.com/resources/news/2012/08/30/9584929/tighter-sanctions-fail-to-deter-iran-methanol-exports-to-asia/ retrieved on March 17, 2023.

3 | METHODOLOGY

This section presents the empirical methodology. To begin with, we briefly review the behaviour of a price-taking exporter that directs its shipments to a collection of destination markets and uses it to derive a condition for spatial price integration between two markets. We then show how an empirical strategy based on the estimation of a parity bound model can be used to detect any possible violation of that condition and measure the probability of observing departures from that condition. Lastly, we review how the extended PBM in Negassa and Myers (2007) provides an adapted testing strategy to investigate whether the sanctions have affected the degree of spatial integration.

3.1 | Theoretical predictions and empirical strategy

Following Ritz (2014), we consider a methanol producer k that sells a positive output to $M \geq 2$ export markets. In each market i, the excess demand in period t is given by the smooth and strictly decreasing function $q_{it} = D\left(P_{it}\right)$ where q_{it} and $P_{it,}$ respectively, represent the total imports and the local market clearing price. We let s_{it}^{k} denote the quantity shipped by k to market i.

The producer's processing cost $C_k\left(\sum_{i=1}^{M} s_{it}^{k}\right)$ is given by a well-behaved function of the total quantities sold in all export markets. In addition, the producer k incurs a market-specific transport cost τ_{it}^{k} per unit of output sold in region i in period t. This mainly reflects the cost of shipping and may vary across destination markets depending on distance and other factors. Production is also subject to a capacity constraint K_k such that $\sum_{i=1}^{M} s_{it}^{k} \leq K_k$.

Assuming perfect competition (thus a price-taking behaviour),[15] producer k's profit-maximisation problem in period t is to decide the amount of methanol to export to each market:

$$\underset{s_{it}^{k}}{Max} \prod_{k}\left(s_{it}^{k}\right) = \sum_{i=1}^{M} P_{it} s_{it}^{k} - C_k\left(\sum_{i=1}^{M} s_{it}^{k}\right) - \sum_{i=1}^{M} \tau_{it}^{k} s_{it}^{k} \tag{1}$$

$$s.t. \quad \sum_{i=1}^{M} s_{it}^{k} \leq K_k$$

The Lagrangian for that constrained optimisation problem is $\Pi_k\left(s_{it}^{k}\right) + \zeta_k\left(K_k - \sum_{i=1}^{M} s_{it}^{k}\right)$ where ζ_k is the non-negative shadow value of the capacity constraint. We consider a given pair of destination markets (i,j) and assume that this problem has an interior solution for these two export markets. Denoting by MC_k the producer's marginal cost of production, the first-order conditions of optimality for these two markets are such that, in each market, the marginal revenue (i.e. the local market-clearing price) equals the sum of the marginal production cost, the unit transportation cost and the shadow value of the capacity constraint, that is,

[15]On the supply side, the market structure exhibits a limited degree of concentration as the 10 largest methanol processing firms solely control 26% of the world production capacity (IHS, 2017). On the demand side, a fragmented demand structure also prevails as this commodity is utilised in different industrial sectors to produce a range of different products (IHS, 2021). Within a country, the demand for methanol typically emanates from a collection of small independent firms with low market concentration ratios (see e.g. CAAFEA, 2021; Tang et al., 2009 for a discussion on the case of China).

$P_{it} - MC_k - \tau_{it}^k - \zeta_k = 0$ and $P_{jt} - MC_k - \tau_{jt}^k - \zeta_k = 0$. These conditions jointly indicate that the spatial price spread verifies:

$$P_{it} - P_{jt} = \tau_{it}^k - \tau_{jt}^k. \tag{2}$$

In that equation, the observed spatial price differential between the two markets is entirely determined by the difference in unit transportation costs at time t. Neither the producer's marginal production cost nor the shadow value of the capacity constraint intervenes in the determination of the observed spatial price gap. That relation shows that the behaviour of the supplier contributes to the economic integration of the two markets as the local prices verify Marshall's law of one price. Hereafter, we attempt to verify whether that equation empirically holds or not. That equation explains why we concentrate on price differences rather than on price levels.

We let T_t denote the difference in unit transportation costs at time t. Following Sexton et al. (1991), we consider a taxonomy of three mutually exclusive trade regimes governing the spatial price spreads observed between i and j.

In Regime I, the spatial price spread is equal to the unit transportation cost difference and is thus said to be 'at the parity bounds':

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

By construction, this regime is consistent with the conditions for the profit maximisation of a producer that supplies the two markets. As highlighted in Barrett and Li (2002), this regime is consistent with the conditions for perfect spatial market integration.

In Regime II, the spatial price difference is greater than the unit cost difference and hence 'outside the parity bounds':

$$P_{it} - P_{jt} - T_t > 0. \tag{4}$$

In this regime, there is either a relative shortage situation in market i at time t in that less product was allocated to that destination than indicated by the efficient arbitrage condition (3) or a market glut in market i. Anyway, markets are separated and there are unseized opportunities for a profitable rebalancing of the supplier's exports that consist of shifting the quantities initially destined to market j in the direction of market i. This regime may conceivably result from a host of factors including: import restrictions in market i, capacity constraints at port infrastructures or shipping scarcities caused by the unavailability of chemical tankers.

In Regime III, the spatial price difference is strictly lower than the unit cost difference and hence 'inside the parity bounds':

$$P_{it} - P_{jt} - T_t < 0. \tag{5}$$

This regime is the opposite of the previous one: There are unseized opportunities for a profitable rebalancing of the export flows from market i in the direction of j.

3.2 | An adapted parity bounds model

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. We first present the standard PBM in Sexton et al. (1991) before reviewing the extended version proposed in Negassa and Myers (2007).

3.2.1 | The standard PBM

From an empirical perspective, the unit cost difference T_t is not available to the modeller. In the present study, it was not possible to obtain quotations for the time charter rates and the destination-specific insurance costs incurred for oceangoing chemical tankers. So, we follow Sexton et al. (1991) and assume that the unit cost difference T_t can be modelled as follows:

$$T_t = \alpha + Z_t\beta + \varepsilon_t, \tag{6}$$

where α and β are unknown parameters to be estimated, Z_t is a vector of explanatory variables aimed at controlling for exogenous effects (e.g. seasonal dummy variables) and ε_t is a random error that is assumed to be i.i.d. normally distributed with a zero mean and a standard deviation σ_ε. By construction, that specification captures the time-invariant and the seasonal components of that unit cost difference.

Using that specification, the three regimes can be written:

$$\text{Regime I:} \quad P_{it} - P_{jt} = \alpha + Z_t\beta + \varepsilon_t, \tag{7}$$

$$\text{Regime II:} \quad P_{it} - P_{jt} = \alpha + Z_t\beta + \varepsilon_t + \eta_t, \tag{8}$$

$$\text{Regime III:} \quad P_{it} - P_{jt} = \alpha + Z_t\beta + \varepsilon_t - \eta_t, \tag{9}$$

where η_t is a one-sided, positive half-normal error that is independent of ε_t and has a standard deviation parameter σ_η.

For concision, we let $\pi_t = P_{it} - P_{jt} - \alpha - Z_t\beta$ denote the random variable that gives the expected value of the difference between the spatial price spread and the deterministic component of the unit cost difference T_t at time t. The distribution functions f_t^I, f_t^{II} and f_t^{III} for this random variable in each of the three regimes are detailed in Table 1.

The standard PBM of Sexton et al. (1991) consists of a switching regression model with three regimes that posits that the probabilities λ_I, λ_{II} and $\lambda_{III} = 1 - \lambda_I - \lambda_{II}$ of observing each regime are constant over time. The joint density function for π_t over all trading regimes is then given as the mixture distribution:

$$f_t(\pi_t) = \lambda_I f_t^I(\pi_t) + \lambda_{II} f_t^{II}(\pi_t) + [1 - \lambda_I - \lambda_{II}] f_t^{III}(\pi_t) \tag{10}$$

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

$$L = \prod_{t=1}^{N} f_t(\pi_t). \tag{11}$$

The model parameters—namely the mean parameters α and β, the standard deviations σ_ε and σ_η and the regime probabilities λ_I and λ_{II}—can be estimated by maximising the logarithm of this

TABLE 1 The density functions of the three regimes.

Regime I: $f_t^I(\pi_t) = \frac{1}{\sigma_\varepsilon}\phi\left(\frac{\pi_t}{\sigma_\varepsilon}\right)$
Regime II: $f_t^{II}(\pi_t) = \left[\frac{2}{\sqrt{\sigma_\varepsilon^2+\sigma_\eta^2}}\right]\phi\left(\frac{\pi_t}{\sqrt{\sigma_\varepsilon^2+\sigma_\eta^2}}\right)\left[1-\Phi\left(\frac{-\pi_t\sigma_\eta}{\sigma_\varepsilon\sqrt{\sigma_\varepsilon^2+\sigma_\eta^2}}\right)\right]$
Regime III: $f_t^{III}(\pi_t) = \left[\frac{2}{\sqrt{\sigma_\varepsilon^2+\sigma_\eta^2}}\right]\phi\left(\frac{\pi_t}{\sqrt{\sigma_\varepsilon^2+\sigma_\eta^2}}\right)\left[1-\Phi\left(\frac{\pi_t\sigma_\eta}{\sigma_\varepsilon\sqrt{\sigma_\varepsilon^2+\sigma_\eta^2}}\right)\right]$

Note: Here, ϕ denotes the standard normal density function, and Φ is the standard normal cumulative distribution function. The density function of Regime I is that of a normal variable. Those of regimes II and III are the density of the sum of a normal random variable and a truncated normal random variable derived in Weinstein (1964).

likelihood function subject to constraints that impose that the regime probabilities lie between zero and one and that the standard deviations are non-negative.[16]

Once the model has been estimated, one can use the estimation results to explore whether the regime probabilities change over time. Following Kiefer (1980) or Spiller and Wood (1988, p.889–90), we can evaluate the probability Proba_t^r that the observation at time t was generated by regime r given the estimated values of the model parameters. This probability is: $\text{Proba}_t^r = \hat{\lambda}_r f_t^r(\pi_t) / \left[\hat{\lambda}_I f_t^I(\pi_t) + \hat{\lambda}_{II} f_t^{II}(\pi_t) + \left[1-\hat{\lambda}_I-\hat{\lambda}_{II}\right] f_t^{III}(\pi_t)\right]$. A visual inspection of the time series Proba_t^r can provide useful insights on whether a regime was clustered during some specific period (e.g. during the sanctions). That said, in order to statistically investigate whether our assumption of time-invariant regime probabilities λ_I, λ_{II} and $\lambda_{III} = 1 - \lambda_I - \lambda_{II}$ is or is not supported by the data, one has to consider the following more general specification.

3.2.2 | An extended PBM

We now briefly review the extended PBM specification proposed in Negassa and Myers (2007) and relax the assumption of time-invariant regime probabilities. Here, we want to explore whether the sanctions have modified the probabilities of observing the various regimes. So, we allow the regime probabilities to change when the sanctions are imposed.

We consider two periods within the sample depending on whether international sanctions were imposed on Iranian exports or not. In each period, the trading regime probabilities are constant over time but the probabilities of the two periods can differ.[17] Formally, we now let $(\lambda_I, \lambda_{II}, \lambda_{III} = 1 - \lambda_I - \lambda_{II})$ and $(\delta_I, \delta_{II}, \delta_{III} = 1 - \delta_I - \delta_{II})$ denote the estimated probability of observing regime r absent and under the sanctions, respectively, and D_t denote a dummy variable that takes a value of 1 when sanctions are imposed. Hence, the probability of being in regime r at time t is now $\lambda_r(1 - D_t) + \delta_r D_t$.

[16] A hill-climbing procedure is typically used to numerically solve this constrained optimisation problem. In this paper, all the estimates have been obtained using the Davidson–Fletcher–Powell routine.

[17] In the application discussed 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 because we believe that it is adapted to model the abrupt and immediate nature of the sanctions. Technically, it is possible to relax this assumption by modelling D_t as a transition variable and allowing an intermediary adjustment period during which its value linearly increases from 0 to 1.

The likelihood function for a sample of observations is: $L = \prod_{t=1}^{N} f_t(\pi_t)$ where the joint density function for the observation at time t is now:

$$\begin{aligned} f_t(\pi_t) = {} & \left[\lambda_I(1-D_t)+\delta_I D_t\right] f_t^I(\pi_t) \\ & + \left[\lambda_{II}(1-D_t)+\delta_{II} D_t\right] f_t^{II}(\pi_t) \\ & + \left[(1-\lambda_I-\lambda_{II})(1-D_t)+(1-\delta_I-\delta_{II})D_t\right] f_t^{III}(\pi_t) \end{aligned} \tag{12}$$

Again, the model is estimated by maximising the log-likelihood function subject to constraints stipulating that the regime probabilities have to lie in the unit interval in each period (i.e. $0 \le \lambda_r \le 1$ and $0 \le \delta_r \le 1$ for any regime) and the standard deviations are non-negative.

By construction, this extended specification allows us to test the null hypothesis of no structural change in any regime probability. This can be done using a likelihood-ratio test comparing two nested specifications: the unrestricted one corresponding to that extended PBM model with the restricted one positing no structural change in the regime probabilities (i.e. the probability of observing each regime is kept constant over time: $\lambda_r = \delta_r$).

4 | DATA

The model was applied to the major Asian methanol importing markets to estimate the effects of the trade sanctions on spatial market efficiency. We use monthly[18] transaction price data for methanol delivered in China, India, South Korea and South-Eastern Asia (which includes Indonesia, Malaysia, Philippines, Singapore, Thailand and Viet Nam) issued by Argus, a price-reporting service (Argus, 2021).[19] Altogether, these countries accounted for 66% of the global consumption of methanol in 2017 (IHS, 2017). These prices are denominated in US dollars per metric ton. These transaction prices refer to a common incoterm (CFR) and concern comparable transactions lots (i.e. the standard cargo of an oceangoing chemical vessel). We do not consider the US market because the importation of Iranian petrochemical products never occurred during the period under scrutiny (it was banned under US unilateral sanctions). Similarly, we do not consider Europe because that market area exhibits important institutional differences that make it barely comparable to the Asian ones.[20]

We consider the period covering January 2009 to October 2018 which yields a sample size of 118 observations. The sample starting date is imposed by the unavailability of methanol price data in India prior to 2009. The end of the sample period is the last price quotation before the re-imposition of the US sanctions against Iran. Within that sample period, we further distinguish

[18] This data frequency is appropriate in our context because one can expect a methanol exporter to be able to respond within 30 days to the profit opportunities that may emerge at the destination markets.

[19] Our methodology and data use price data at the market level, not only for the methanol coming from Iran. The standard parity bounds empirical model is also based on price information only. Still, Barret and Li (2002) and Massol and Banal-Estanol (2018) use volume data alongside price data to identify whether trade in a given direction occurred or not. Furthermore, in the Argus database, the Southeast Asian economies are treated as a single entity. There is no spot market for methanol delivered in Japan in this database. Because of the absence of market-based price quotations for methanol, we cannot include Japan in the present analysis.

[20] In Europe, methanol is quoted using an FOB incoterm and the transactions concern smaller units (i.e. the standard lot traded in Europe concerns a regional barge delivery from Rotterdam).

TABLE 2 Average prices at destination (in $/ton).

	China	India	SE Asia	S. Korea
Entire sample period				
Mean price	324.46	299.78	339.07	339.12
Subperiod I: before the sanctions				
Mean price	309.49	280.13	303.53	307.88
Subperiod II: under the sanctions				
Mean price	345.29	321.28	376.50	369.19
Variation relative to subperiod I	*+11.6%*	*+14.7%*	*+24.0%*	*+19.9%*
Subperiod III: after the sanctions				
Mean price	314.19	294.30	331.10	335.97
Variation relative to subperiod I	*+1.5%*	*+5.1%*	*+9.1%*	*+9.1%*

the subperiod from May 2012 to January 2016 (i.e. 45 observations) during which Iran's petrochemical exports were subjected to international sanctions.

Although the analysis below is centred on the analysis of the spatial price spreads between two markets, a few preliminary remarks can be made on the average prices observed at these four markets. Table 2 indicates that, during the whole sample period, methanol was cheaper on average in India and China than in the other two regions. Unsurprisingly, the average prices observed when the sanctions are imposed are noticeably larger than either before or after that period. It is interesting to stress that the magnitude of the average price increases observed under the sanctions is not uniform across all markets. In Korea and in Southeast Asia, the average price is, respectively, 19.9% and 24% larger than that observed during the preceding period whereas in China and in India that increases is more modest (respectively, 11% and 14.6%).

The last remark calls for further investigations on the evolution of the spatial price spreads. Figure 1 provides plots of these six price differentials (in levels) that can be formed using the four price series. A visual inspection of these plots suggests that the price gap between South Korea and Southeast Asia is less variable than that of other market pairs (e.g. Southeast Asia and India). One can also note that, for some pairs (e.g. Korea–China, Korea–India), the variability of the observed spatial price spread seems to be larger when the sanctions are implemented.

Table A1 (in Appendix A) summarises the statistical properties of these six price differentials. These price gaps are modestly skewed and occasionally fat-tailed (e.g. China–India, Korea–Southeast Asia). The first-order autocorrelation coefficients reveal evidence of positive serial correlation for all series, which is consistent with the slow-moving nature of the logistics involved in that trade. An oceangoing chemical tanker is a slow-moving transportation system and its redirection can typically extend the time spent at sea by 1 or 2 weeks. Hence, the price differential observed at time t is likely to represent the outcome of decisions taken both during the current observation and the preceding one. As an unmodelled autocorrelation can result in inefficient estimates, we correct for the presence of serial correlation using the modified specification detailed in Appendix B.

We examine the time-series properties of these price gaps by testing for the presence of a unit root over the full sample. We use three standard unit root tests, namely the Augmented Dickey Fuller test, the Phillips–Perron test and the modified version of the Phillips–Perron test proposed by Ng and Perron (2001), to examine whether the null hypothesis of a unit root can be rejected in favour of the

FIGURE 1 Data plots of the price differential series (USD/ton).

alternative of either mean stationarity or trend stationarity. The results are reported in Table A2. In this paper, we use a test size equal to 10% and hereafter consider all the price differentials such that the null hypotheses are rejected at that level.[21] Among these six price gap series, we find that four (i.e. China–India, Southeast Asia–India, Korea–India and Korea–Southeast Asia) are mean stationary and that one (namely, Korea–China) is trend stationary. This last finding is not surprising as the plot for the Korea–China pair in Figure 1 also exhibits an upward trend. We thus include a trend in the

[21]By choosing a large test size, we prefer to err on the side of declaring a price differential stationary and using that series to estimate the parity bound model. The 10% threshold value is also used in Cuddington and Wang (2006) to identify stationary series.

specification used for that specific market pair. In contrast, the price differential between China and Southeast Asia is not mean reverting as the two tests 'fail to reject' the unit root hypothesis at the 10% level. As the estimation of our statistical models to the China–Southeast Asia price differential conveys the risk of generating spurious results, that price spread will not be considered in the sequel.

In the analysis below, some specifications include a list of stationary control variables: (i) the monthly industrial production growth rates reported by the Asian Development Bank and (ii) the percentage change in the monthly price of crude oil that was computed from the crude oil price series reported by the World Bank. The industrial production series are aimed at controlling for local, market-specific demand shocks. Variations in the price of crude oil can affect both the demand for methanol (as it can be consumed as a fuel or as a gasoline additive) and the shipping cost of methanol (because of its impact on the cost of the heavy fuel consumed by chemical tankers).

5 | EMPIRICAL RESULTS AND DISCUSSION

We first present the results obtained estimating standard linear regression models and then the ones obtained using the PBM framework.

5.1 | Preliminary insights from linear regressions

A visual inspection of Figure 1 suggests that, for some market pairs, the price differentials could be widened when the sanctions are imposed. To further investigate that, we conduct a series of preliminary analyses and estimate, for each price gap series that is stationary, a linear regression of the form: $P_{it} - P_{jt} = \alpha + Z_t\beta + D_t\gamma + \varepsilon_t$ where: D_t is the aforementioned dummy variable signalling the sanctions; Z_t is a vector of observable control variables that includes a trend whenever the price gap series is trend stationary; α, β and γ are the coefficients to be estimated; and ε_t is the error term. We consider two versions of that specification. In the first one (labelled Model I), the vector of exogenous regressors Z_t includes 11 monthly variables to control for possible seasonal effects. In the second one (labelled Model II), that vector includes: (i) the industrial production growth rates in markets i and j, (ii) the percentage change in the monthly price of crude oil and (iii) three quarterly dummy variables to control[22] for possible seasonal effects. Because of the autocorrelated nature of the price gap series, we also include an AR(p) structure to correct for autocorrelation.

The estimation results are presented in Table 3. The estimates obtained with the two linear models are similar. First, in both cases, the sanction coefficient is not statistically significant for the pairs China–India and Korea–Southeast Asia. This finding indicates that the price spreads observed under the sanctions are not different from those observed absent the sanctions, which suggests that the sanctions have not statistically impacted the law of one price linking the two markets. Second, and in contrast, the estimates of the sanction coefficient obtained under the two models are positive and statistically significant for the market pairs Southeast Asia–India, Korea–China and Korea–India. This means that, for these pairs, there is a widened spread when sanctions are imposed. As these spreads are positive, Southeast Asia and Korea experienced noticeably greater prices than the ones in China and India when Iran's exports were subjected to sanctions.

[22]In this second specification, we allow only for quarterly dummies and not monthly to save degrees of freedom.

TABLE 3 Estimation results for the preliminary investigations based on linear specifications.

	China–India		SE Asia–India		S. Korea–China		S. Korea–India		S. Korea–SE Asia	
	Model I	**Model II**	**Model I**	**Model II**	**Model I**	**Model II**	**Model I**	**Model II**	**Model I**	**Model II**
Constant	32.046***	19.818**	41.561***	31.846***	−5.567	−8.013	44.118***	32.446***	2.297	2.125
	(6.945)	*(8.666)*	*(7.131)*	*(5.635)*	*(5.637)*	*(9.142)*	*(5.534)*	*(4.259)*	*(5.094)*	*(4.631)*
Trend	—	—	—	—	0.341***	0.329***	—	—	—	—
					(0.069)	*(0.079)*				
$Sanctions_t$	−2.395	1.382	18.241***	22.696***	9.823**	10.952**	11.949***	15.137***	−6.266	−5.039
	(5.984)	*(5.817)*	*(6.939)*	*(6.210)*	*(4.302)*	*(4.285)*	*(4.332)*	*(4.390)*	*(4.701)*	*(4.570)*
$Crude\_Var_t$	—	0.304*	—	0.208	—	−0.062	—	0.216	—	−0.041
		(0.175)		*(0.166)*		*(0.100)*		*(0.145)*		*(0.070)*
Ind_Prod_{it}	—	0.426	—	−0.265	—	0.393	—	0.221	—	−0.050
		(0.703)		*(0.422)*		*(0.254)*		*(0.351)*		*(0.170)*
Ind_Prod_{jt}	—	0.249	—	0.871**	—	−0.102	—	0.518	—	−0.101
		(0.428)		*(0.421)*		*(0.521)*		*(0.353)*		*(0.208)*
Seasonal dummies	Monthly	Quarterly	Monthly	Quarterly	Monthly	Quarterly	Monthly	Quarterly	Monthly	Quarterly
AR(p) lag structure	1,2	1,2	1,2	1,2	1	1	1,2	1,2	1,2,3	1,2,3
Adjusted R^2	.363	.394	.596	.621	.759	.759	.450	.471	.682	.684
Diagnostic tests										
L-B(6)	3.695	3.961	3.360	4.212	6.390	5.947	1.204	1.646	4.638	3.936
ARCH(1)	0.439	0.144	0.541	0.614	0.019	0.266	3.342*	1.093	0.311	0.368

Note: Estimates for the monthly dummies and the AR(p) coefficients are not reported for brevity. Numbers in parentheses are standard errors. Asterisks indicate significance at .1*, .05** and .01*** levels respectively. As discussed in Section 4.1, the specification used for the pair (Korea, China) is supplemented with a trend because that explained variable is trend stationary. L-B(6) is the Ljung-Box Q-statistic for the null hypothesis of no serial correlation up to the sixth order. ARCH(1) is the LM-test for the absence of autoregressive conditional heteroscedasticity with 1 lag.

Overall, these preliminary investigations suggest a change in the market fundamentals affecting the formation of the spatial price gaps in that multimarket system. Given the profound trade reconfigurations resulting from the sanctions, one can suspect that occasional disruptions could have affected the spatial arbitrages performed between these spatially distinct markets. As a variety of factors (e.g. episodic transportation bottlenecks) can affect the tradability of methanol between these destinations, the spatial arbitrages linking these destinations markets can exhibit discontinuities. As such discontinuities cannot be appropriately captured using a simple linear specification with time-invariant parameters (see the discussions in Barrett, 1996, 2001; Baulch, 1997; McNew & Fackler, 1997), we now present the results obtained with the PBM approach.

5.2 | Estimation results using the PBM

To gain insights on the adequacy of a specification positing time-invariant regime probabilities, we first estimate the simplest specification—namely the standard parity bound model with time-invariant regime probabilities—and use the estimated values to generate the probabilities Proba_t^r that the observation at time t was generated by regime r, as described in (7)–(9). Overall, we find that these series show, for any regime, sign of a changing behaviour when the sanctions are imposed. As an illustration, Figure 2 depicts the 15-month centred moving average of these probabilities for two price differentials: China–India and Korea–China. A visual inspection of these plots reveals that, for Korea–China, the probabilities shifted noticeably over time and that the largest shifts occur when sanctions are imposed against Iran. In contrast, the changes in Proba_t^r are less pronounced for China–India. Overall, these observations call for further statistical investigations based on the extended specification.

For each of the five stationary price differentials, we then estimate the extended parity bound model that allows for a possible change in the regime probabilities when sanctions are imposed. As in the preliminary analysis above, we successively consider two definitions for the list of exogenous regressors in Z_t: either 11 monthly variables (Specification I) or an extended list of variables to control for variations in industrial production at destination markets, the relative change in the price of crude oil and three quarterly dummy variables. A trend is also added whenever the price gap series under scrutiny is trend stationary.

The log-likelihoods of the converged solutions obtained with the standard and the extended models are reported in Table 4. That table also reports the likelihood ratio (LR) test for the null hypothesis of zero probability changes during the sanctions.

The tests conducted with the two PBM specifications yield consistent results. Three interesting findings emerge. First, China–India is the only market pair for which the LR tests fail to reject the validity of the restricted model with unchanged probabilities at the 10% level, which reveals that the sanctions have not substantially modified the degree of market integration between these two importing countries. This finding is consistent with the market commentaries that highlight that, in these two countries, the importation of Iranian methanol did not cease under the sanctions.

Second, in contrast, the LR tests firmly reject the null hypothesis at the 1% level for all the market pairs involving either China or India with either Korea or Southeast Asia. The magnitude of the LR test statistic is very high (i.e. larger than 16) which clearly confirms that sanctions affected the arbitrage activities between these market pairs. While South Korea is one of Iran's major trading partners, along with China and India, the finding that sanctions had a different

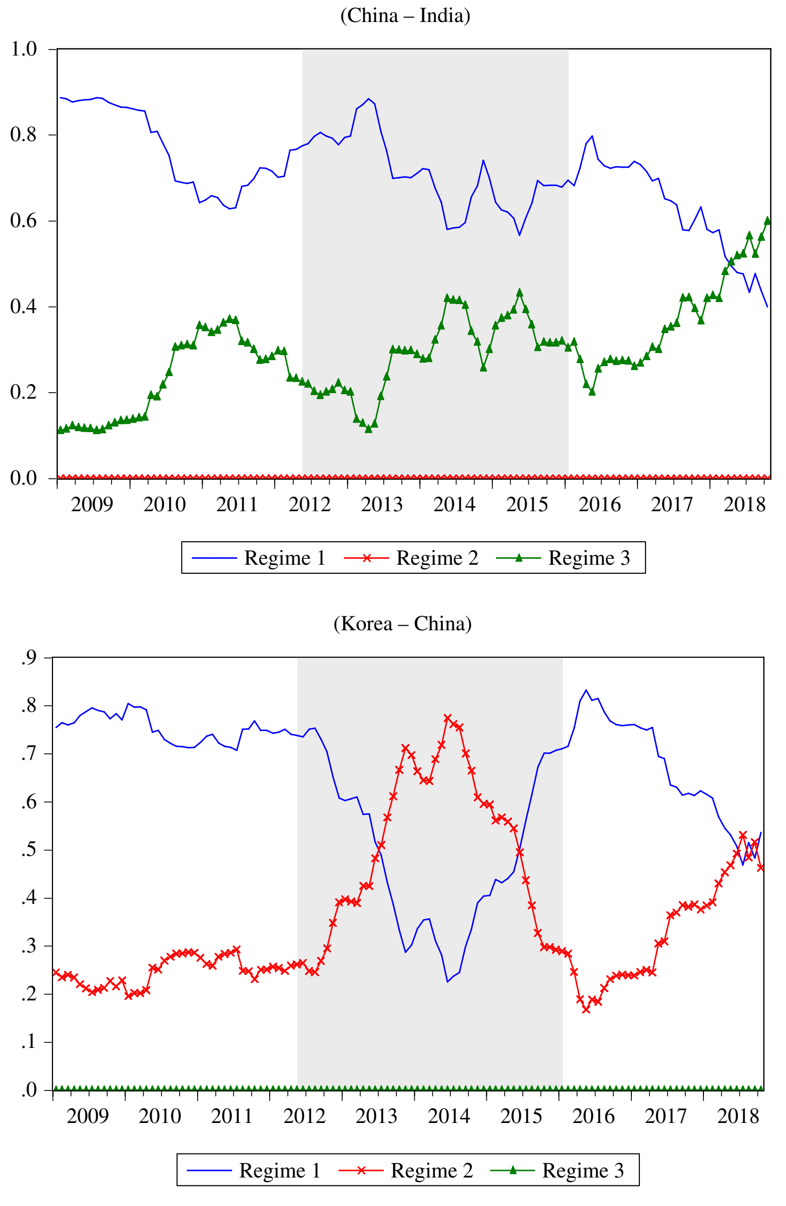

FIGURE 2 Fifteen month-centred moving average estimates of regime probabilities for China–India and Korea–China.

impact on that country is not that surprising because Seoul is reputed to have scaled back its bilateral trade with Iran under the sanctions (CRS, 2021, p. 46).

Lastly, the LR test statistics obtained for the pair Korea–Southeast Asia are substantially lower than the ones obtained with the pairs associating these markets with either China or India. This suggests that the null hypothesis of no change is only mildly rejected for this pair which involves two affected markets. This is not surprising as the arbitrage activity performed between this market pair is also reputed to be performed by other suppliers located in the Middle East (i.e. Saudi Arabia, Bahrain, and Oman) that have not been sanctioned.

TABLE 4 Likelihood ratio tests.

		Log-likelihood		LR test	
		Restricted model	Unrestricted model	$\chi^2(2)$ statistics	(*p*-value)
China–India	Specification I	−474.802	−473.988	1.628	(*.443*)
	Specification II	−478.690	−478.676	0.028	(*.986*)
SE Asia–India	Specification I	−492.797	−477.619	30.356***	(*.000*)
	Specification II	−492.643	−476.103	33.079***	(*.000*)
S. Korea–China	Specification I	−426.280	−416.020	20.520***	(*.000*)
	Specification II	−426.550	−416.684	19.733***	(*.000*)
S. Korea–India	Specification I	−468.940	−460.909	16.063***	(*.000*)
	Specification II	−472.289	−463.417	17.745***	(*.000*)
S. Korea–SE Asia	Specification I	−411.784	−408.880	5.808*	(*.055*)
	Specification II	−414.937	−410.713	8.448**	(*.015*)

Note: Asterisks indicate rejection of the null hypothesis at the .1*, .05** and .01*** significance levels respectively.

We now examine the estimation results. As the LR test fails to reject the hypothesis of unchanged probabilities for the pair China–India, Table 5 reports the estimation results obtained with that restricted model. Table 6 reports the estimation results for the unrestricted model for all the other market pairs. In these tables, we report the estimates obtained for: the transportation cost coefficients, the autocorrelation parameter, the standard deviation coefficients and the regime probabilities obtained absent (λ's) or under the sanctions (δ's). For concision, the seasonal parameters included in the mean equation (i.e. the coefficients of the 11 monthly dummy variables) are not reported. As an illustration, we detail an interpretation of these estimated coefficients before presenting the main findings. In the last column in Table 6, outside of the sanctions period, the probability of that spatial price spread being at the parity bounds is 66.7% whereas that of being inside is 13% and that of being outside is 20.3% (λ coefficients). Hence, there is a high probability (66.7%) that this spatial price difference equals the transport cost difference. Still, there is some probability that the price in South Korea is well above or well below that of Southeast Asia (20.3% and 13%, respectively).

To begin with, note that 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 price gap series. This finding confirms the need to explicitly model serial correlation in our specification.

The estimates in Table 6 show that the results obtained with the two specifications are very similar. Outside of the sanction period, a fairly high degree of economic integration is observed among all the market pairs, as the estimates for λ_I are larger than 58%. The estimated probability even attains 100% for two market pairs, namely Southeast Asia–India and Korea–India. These high estimates are consistent with the market commentaries that typically highlight that a high degree of integration is achieved among the Asian markets when there are no sanctions (e.g. IHS, 2017).

During the sanction period, we observe that the probability of observing Regime I remains significant and large (more than 39%) for the pair Korea–Southeast Asia. This finding shows that, despite its probability being reduced, some level of market integration was preserved under the sanctions. This finding is consistent with the mild rejection of the restricted model in Table 4. In contrast, during the sanctions, we observe a complete reconfiguration of the

TABLE 5 PBM estimation results for the price differential between China and India.

	China–India	
	Specification I	**Specification II**
Mean parameters		
α	44.126***	34.781***
	(4.922)	*(7.763)*
$Crude\_Oil\_var_t$	—	0.193
		(0.150)
Ind_Prod_{it}	—	0.191
		(0.463)
Ind_Prod_{jt}	—	−0.005
		(0.317)
ρ	0.565***	0.541***
	(0.061)	*(0.080)*
Std. deviations		
σ_ε	8.701***	9.561***
	(1.033)	*(1.485)*
σ_η	26.599***	23.083***
	(4.767)	*(4.139)*
Probabilities (in %)		
λ_I	70.762***	66.925***
	(10.659)	*(14.924)*
λ_{II}	0.000	0.000
	(0.031)	—
$1-\lambda_I-\lambda_{II}$	29.238***	33.075***
	(10.659)	*(14.924)*
Log likelihood	−474.802	−478.690

Note: Estimates for the seasonal (i.e. monthly for Specification I and quarterly for Specification II) dummies are not reported for brevity. Numbers in parentheses are standard errors. Significance tests are based on asymptotic standard errors that have been computed using the Hessian matrix of the log-likelihood function. Asterisks indicate significance at .1*, .05** and .01*** levels respectively. A dash — for the standard error indicates not calculated due to the probability estimate being at the boundary of the parameter space.

degree of market integration among either China or India and either Korea or South East Asia. For these three market pairs, Southeast Asia–India, Korea–China and Korea–India, the probability of observing Regime I radically drops to zero when the sanctions are imposed and those of Regime II becomes very high and larger than 90%. Recall that this last regime is such that the price spread observed between Southeast Asia (respectively, Korea) and India (respectively, China) is 'outside the parity bounds' which suggests that the prices in Korea and Southeast Asia are substantially larger than the sum of the ones observed in India and China plus the differences in the unit transportation costs when the sanctions are imposed. This last finding can be interpreted as objective signs of export deflections from Iran to India and China.

TABLE 6 Extended PBM estimation results for the price differentials between the other markets.

	SE Asia–India		S. Korea–China		S. Korea–India		S. Korea–SE Asia	
	Spec. I	**Spec. II**	**Spec. I**	**Spec. II**	**Spec. I**	**Spec. II**	**Spec. I**	**Spec. II**
Mean parameters								
α	38.730***	29.162***	−10.506***	−21.409***	45.515***	34.206***	2.129	2.830
	(5.300)	*(4.033)*	*(3.737)*	*(7.123)*	*(5.355)*	*(3.858)*	*(4.166)*	*(3.626)*
$Crude\_Oil\_var_t$	—	0.251*	—	−0.106	—	0.247*	—	0.000
		(0.141)		*(0.093)*		*(0.141)*		*(0.086)*
Ind_Prod_{it}	—	−0.102	—	−0.028	—	−0.040	—	−0.074
		(0.323)		*(0.202)*		*(0.338)*		*(0.172)*
Ind_Prod_{jt}	—	0.845**	—	0.745*	—	0.462	—	−0.070
		(0.381)		*(0.434)*		*(0.356)*		*(0.177)*
Trend	—	—	0.274***	0.350***	—	—	—	—
			(0.035)	*(0.051)*				
ρ	0.667***	0.704***	0.581***	0.587***	0.585***	0.614***	0.716***	0.729***
	(0.057)	*(0.060)*	*(0.056)*	*(0.060)*	*(0.068)*	*(0.068)*	*(0.049)*	*(0.045)*
Std. deviations								
σ_ε	11.856***	12.149***	5.503***	5.534***	10.658***	11.456***	3.987***	4.544***
	(0.947)	*(0.922)*	*(0.770)*	*(0.783)*	*(1.220)*	*(1.146)*	*(0.966)*	*(0.804)*
σ_η	20.013***	20.014***	13.272***	13.615***	13.169**	12.886***	12.603***	12.500***
	(3.104)	*(3.173)*	*(1.690)*	*(1.757)*	*(6.178)*	*(4.407)*	*(1.727)*	*(1.825)*
Probabilities (in %)								
λ_I	100.000‡	100.000‡	75.659***	76.470***	100.000‡	99.918***	58.705***	66.714***
	—	—	*(11.446)*	*(10.741)*	—	*(2.082)*	*(16.102)*	*(16.506)*
λ_{II}	0.000	0.000	24.341**	23.530**	0.000	0.000	27.925*	20.332
	—	—	*(11.446)*	*(10.740)*	—	—	*(15.284)*	*(15.840)*

(Continues)

TABLE 6 (Continued)

	SE Asia–India		S. Korea–China		S. Korea–India		S. Korea–SE Asia	
	Spec. I	**Spec. II**	**Spec. I**	**Spec. II**	**Spec. I**	**Spec. II**	**Spec. I**	**Spec. II**
$1-\lambda_I-\lambda_{II}$	0.000	0.000	0.000	0.000	0.000	0.082	13.369*	12.954*
	—	—	*(0.047)*	*(0.069)*	—	*(2.082)*	*(7.203)*	*(7.466)*
Probabilities (in %)								
δ_I	0.008	0.001	0.000	0.002	0.719	0.356	50.556***	39.941*
	(0.451)	*(0.130)*	—	*(0.259)*	*(28.651)*	*(8.525)*	*(17.384)*	*(20.787)*
δ_{II}	95.978***	99.998***	100.000‡	99.998***	89.677***	97.517***	0.000	0.000
	(4.370)	*(0.178)*	—	*(0.265)*	*(28.853)*	*(11.615)*	—	—
$1-\delta_I-\delta_{II}$	4.014	0.000	0.000	0.000	9.604	2.128	49.444***	60.059***
	(4.350)	—	—	*(0.025)*	*(9.591)*	*(7.616)*	*(17.384)*	*(20.787)*
Log likelihood	−477.619	−476.103	−416.020	−416.684	−460.909	−463.417	−408.880	−410.713

Note: Estimates for the seasonal (i.e. monthly for Specification I and quarterly for Specification II) dummies are not reported for brevity. Numbers in parentheses are standard errors. Significance tests are based on asymptotic standard errors that have been computed using the Hessian matrix of the log-likelihood function. Asterisks indicate significance at .1*, .05** and .01*** levels respectively. The symbol ‡ signals that the value is at the upper bound of the probability space. A dash — for the standard error indicates not calculated due to the parameter estimate being at the boundary of the parameter space. As discussed in Section 4.1, the specification used for the pair (Korea, China) is supplemented with a trend because that explained variable is trend stationary.

6 | CONCLUDING REMARKS

The fundamental public policy issue examined in this paper is whether, in a globalised world, the imposition of trade sanctions decided by a group of countries on the exports of a target country affects the prices formed at third-party importing markets and their degree of market integration. Achieving a high degree of market integration is crucial for a smooth allocation of trade flows from surplus to deficit areas, for allowing production decisions to be made according to a comparative advantage, and thus, for the welfare gains from trade to materialise. Incidentally, our analysis also provides an assessment of the effectiveness of the sanctions as it allows us to detect whether, and how, the sanctions are circumvented by some of these third-party nations. These are particularly important questions when the coerced nation is a large producer of a commodity that has a global character.

By means of an extended parity bounds model, we investigate the impacts of the sanctions imposed on the exportation of Iranian petrochemicals using data on the Asian spot markets for methanol during the period 2009–2018. The estimation results provide a series of original findings. Outside of the sanction period, a high degree of market integration is achieved among Asian markets. In contrast, we observe a complete reconfiguration of the spatial extent of these methanol markets under the sanctions. Methanol markets become more fragmented and form two distinct market areas, respectively, comprising China and India on the one hand and Korea and Southeast Asia on the other hand. The degree of market integration achieved within each of these two areas remains very high as we find a high probability of these market pairs being 'at the parity bound' when sanctions are imposed. In contrast, that probability is very low for the market pairs involving one country in each market area which can be interpreted as objective signs of balkanisation.

These findings corroborate the 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.[23] Overall, these results document the importance of Iran as an exporter with an arbitrage behaviour that contributes to the integration of the Asian methanol markets. In the absence of Iran, methanol markets become fragmented. Our results can also be related to the geopolitical discussions on the strained diplomatic relations between China and Western countries, as they provide an example of the asymmetric impact of the trade sanctions imposed by the latter countries. China was indeed less affected than Korea and Southeast Asia by the sanctions imposed on Iran.

Future possible research directions could include further analysis of monthly trade volumes to corroborate the findings. However, to the best of our knowledge, these data are not available. Should this limitation be loosened in the future, the development of an empirical analysis of these trade flows may inform further international methanol trade issues estimating, for instance, the effects of the sanctions on the producer's revenues as well as on the various market welfare levels. Another possible research avenue could concentrate on the effects of the sanctions on the observed price levels (and not on the observed price spreads) before, during and/or after the sanctions. Such an empirical analysis could, for example, involve the specification of a structural econometric model representing the supply and demand of methanol in these importing regions. Yet, that analysis can be complex because methanol supply results from the processing of an exhaustible resource (natural gas) which would mean that one needs to take into account the geological endowment in

[23]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).

natural gas and the intertemporal effects resulting from its extraction. Lastly, we believe that future research can also apply the framework presented in this paper to examine the effects of the sanctions recently imposed on Russian exports following the 2022 invasion of Ukraine.

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DATA AVAILABILITY STATEMENT

The data that support the findings of this study are available from Argus. Restrictions apply to the availability of these data, which were used under licence for this study. Data are available from the authors with the permission of Argus.

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APPENDIX A

TABLE A1 Descriptive statistics of the price gaps.

	Mean	SD	Min	Max	Sk.	Ku.	JB	AC
China–India	24.634	19.506	−37.250	65.750	−0.913	4.167	23.073***	0.594
China–SE Asia	−14.674	25.095	−82.500	40.000	−0.400	3.169	3.284	0.883
SE Asia–India	39.308	23.273	−2.000	105.000	0.495	2.903	4.860*	0.764
S. Korea–China	14.694	17.154	−20.000	57.500	0.261	2.333	3.520	0.845
S. Korea–India	39.328	17.473	−2.500	74.000	0.022	2.450	1.497	0.639
S. Korea–SE Asia	0.020	13.633	−43.125	32.125	−0.492	4.429	14.798***	0.789

Note: The sample period contains 118 observations. The table reports the mean, standard deviation (SD), minimum (min), maximum (max), skewness (Sk.), kurtosis (Ku.) and first-order serial correlation (AC) statistics of the price differentials in levels. It also reports the Jarque–Bera normality test (JB) where asterisks indicate rejection of the null hypothesis of normality at the .1*, .05** and .01*** significance levels respectively.

TABLE A2 Unit root tests.

	Specification[a]	ADF[b]	PP[c]	MZ_t[d]
China–India	C	−5.679***	−4.609***	−3.881***
China–SE Asia	C	−2.348	−2.404	−1.380
SE Asia–India	C	−3.941***	−3.694***	−3.061***
S. Korea–China	C + T	−4.068***	−3.824**	−3.578***
S. Korea–India	C	−5.029***	−4.842***	−3.966***
S. Korea–SE Asia		−2.646***	−3.523***	−2.079**

Notes: Asterisks indicate rejection of the null hypothesis at the .1*, .05** and .01*** significance levels respectively.
[a]C and T are constant and trend respectively. We follow the general-to-specific approach that consists in first including a trend and a constant and successively dropping them whenever the estimated coefficients are not statistically significant at the 5% level.
[b]ADF is the Augmented Dickey–Fuller test with a number of lags suggested by the Bayesian information criterion (BIC).
[c]PP is the Phillips and Perron test based on the Bartlett kernel with bandwidth selected from the Newey–West method.
[d]MZ_t is the modified Phillips–Perron test in Ng and Perron (2001) with a number of lags suggested by the BIC.

APPENDIX B

To correct for the presence of serial correlation, we extend the model above to account for first-order autocorrelation in the error term using the modified specification:

$$\text{Regime I:} \quad P_{it} - P_{jt} = \alpha + Z_t\beta + \mu_t, \quad \text{where} \quad \mu_t = \rho\mu_{t-1} + \varepsilon_t. \tag{B1}$$

$$\text{Regime II:} \quad P_{it} - P_{jt} = \alpha + Z_t\beta + \mu_t, \quad \text{where} \quad \mu_t = \rho\mu_{t-1} + \varepsilon_t + \eta_t. \tag{B2}$$

$$\text{Regime III:} \quad P_{it} - P_{jt} = \alpha + Z_t\beta + \mu_t \quad \text{where} \quad \mu_t = \rho\mu_{t-1} + \varepsilon_t - \eta_t. \tag{B3}$$

Here, ρ is an autocorrelation parameter to be estimated (with $-1 < \rho < 1$), μ_t (respectively, μ_{t-1}) is the observed current (respectively, lagged) residual representing the difference between the spatial price differential $P_{it} - P_{jt}$ and the deterministic component of the transportation parameter T_t.