Market power across the Channel: Are Continental European gas markets isolated? ☆

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Abstract

This paper examines the efficiency of the arbitrages performed between two regional markets for wholesale natural gas linked by a capacity-constrained pipeline system. We develop a switching regime specification to (i) detect if the observed spatial arbitrages satisfy the integration notion that all arbitrage opportunities between the two markets are being exploited, and (ii) decompose the observed spatial price differences into factors such as transportation costs, transportation bottlenecks, and the oligopolistic behavior of the arbitrageurs. Our framework incorporates a test for the presence of market power and it is thus able to distinguish between the physical and behavioral constraints to marginal cost pricing. We use the case of the "Interconnector" pipeline as an application, linking Belgium and the UK. Our empirical findings show that all the arbitrage opportunities between the two zones are being exploited but confirm the presence of market power.

Keywords: Law of one price, market integration, spatial equilibrium, interconnectors, natural gas.

JEL Classification: D43; F14; F15; L95; Q37.

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1. Introduction

Over the last two decades, a series of structural and regulatory reforms have been carried out to promote a competitive organization for both the North American and European natural gas industries. A significant development in this restructuring was the emergence of spatially localized spot markets for wholesale natural gas that are interconnected throughout the pipeline network. By opening access to the pipeline system, these liberalization reforms have allowed gas arbitrageurs to purchase transportation rights and thus compete to exploit spatial price differences between interconnected markets. The efficiency of these spatial arbitrages represents a major regulatory policy issue. For example, the current European policy debates related to the organization of the EU's internal market for natural gas repeatedly underline the importance of spatial arbitrages as a means to prevent balkanization and ensure an efficient supply of natural gas (Vazquez et al., 2012).

From a theoretical perspective, the definition proposed in Stigler and Sherwin (1985) indicates that two geographical markets for a tradable good are integrated if the spatial price difference between these two markets equals the unit transportation cost. However, from an empirical perspective, assessing the geographic expanse of wholesale gas markets represents a challenging task because intermarket price spreads could reflect a variety of other factors, including transportation bottlenecks and oligopolistic pricing by the arbitrageurs. To overcome this problem, we define integration using the equilibrium notion that all spatial arbitrage opportunities between the two markets are being exploited. This notion is derived from the theoretical literature on spatial price determination that was pioneered by Enke (1951), Samuelson (1952), and Takayama and Judge (1971).

This paper develops an empirical methodology to assess the arbitrages performed between two regional markets for wholesale natural gas linked by a capacity-constrained pipeline system. This methodology is designed to (i) detect if these markets are "integrated," i.e., if all the spatial arbitrage opportunities are being exploited, and (ii) decompose the observed spatial price differences into factors such as transportation costs, transportation bottlenecks, and the oligopolistic behavior of the arbitrageurs. Our framework incorporates a test for the presence of market power and is thus able to distinguish between physical and behavioral constraints to marginal cost pricing. As an application, we use the spatial arbitrages in the "Interconnector" pipeline which connects Europe's two oldest spot markets for natural gas: the UK's National Balancing Point and the Zeebrugge market in Belgium.[1]

A large amount of empirical research has examined the degree of spatial integration between markets for wholesale natural gas with the help of time-series techniques.[2] These studies typically rely

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[1] The application discussed in the paper examines the possible lack of gas flows in the pipeline infrastructure connecting the UK and mainland Europe. That is why the title of this paper is a veiled reference to the British idiom "Fog in the Channel. Continent isolated," allegedly a newspaper headline in Britain in the 1930s announcing the suspension of ferry services between the UK and mainland Europe because of dangerous weather in the English Channel.

[2] A tentative and non-exhaustive methodological clustering of these contributions includes: (i) the early correlation-based analyses (Doane and Spulber, 1994); (ii) the use of Granger causality tests to examine how natural gas price shocks are transmitted across interconnected wholesale markets (Doane and Spulber, 1994); (ii) the pure cointegration-based studies

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 provide useful insights into how local price shocks are transmitted to adjacent markets. However, the methodology used in these studies is of little help in assessing the competitive nature of the observed spatial arbitrages, as they fail to detect the presence of imperfect competition. Moreover, as suggested by the related agricultural economics literature (Barrett, 1996, 2001; Baulch, 1997; McNew and Fackler, 1997), these empirical models are unable to account for the pivotal role played by both intermarket transfer costs and trade flow considerations.[3]

In this paper, we consider an alternative approach based on the parity bounds model (PBM) first introduced in Spiller and Huang (1986).[4] In a PBM, arbitrageurs are assumed to be profit-maximizing agents. Using that assumption, intermarket price spreads are examined using a "switching regime" specification, which estimates the probability of observing each of a series of trade regimes. Sexton et al. (1991), for example, consider three distinct trade regimes: an "arbitrage" regime where the spatial price difference equals the unit intermarket transportation cost, an "autarkic" one where the local prices differ by less than that transportation cost, and a "barriers to trade" regime where the observed spatial price difference is larger than that transportation cost. Barrett and Li (2002), our point of departure, make use of trade flow data to further distinguish whether trade occurs or not in each of the three regimes. This direction-specific approach allows them to detect any violation of the theoretical equilibrium conditions that all arbitrage opportunities between the two markets are being exploited.

We propose a series of modifications of existing PBMs to apply them to the case of natural gas markets. Existing models assume the presence of perfect competition in the spatial arbitrages between two markets. So, we first propose an enriched specification to account for the role of market power and show that it can be used to test for the assumption of perfectly competitive arbitrages. Second, the role of transportation bottlenecks has so far been neglected whereas binding pipeline capacity constraints are likely to occur in the gas industry. So, we propose isolating the specific contribution of pipeline capacity constraints in the observed spatial price spreads. Lastly, the existing PBMs are based on a static formulation where shocks are posited to be serially independent and the variance

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that test for the existence of a long-run equilibrium condition governing the local price series (De Vany and Walls, 1993; Doane and Spulber, 1994; Serletis, 1997; Asche et al., 2002); (iv) the analyses examining a time-varying degree of price convergence among natural gas spot markets with the help of the Kalman Filter approach (King and Cuc, 1996; Neumann et al., 2006; Renou-Maissant, 2012); (v) 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); (vi) the joint assessments of the degree of market integration and price transmission across natural gas markets using tests of cointegration and the corresponding error-correction models (Park et al., 2008; Brown and Yücel, 2008).

[3] *These criticisms emphasize a lack of acquaintance with existing economic models of spatial price determination. Two lines of arguments motivate that shortcoming. First, intermarket transfer costs are typically omitted in these early empirical studies whereas, in theory, price equalizing arbitrage activities are triggered only when localized shocks result in spatial price differences which exceed these intermarket transfer costs (Barrett, 1996, 2001; Baulch, 1997; McNew and Fackler, 1997). Second, trade flows information play no role in these early empirical studies whereas theory suggests that either discontinuities in the trade flows or variations in the directions of these flows can have an impact on the degree of co-movements among prices at each market location (Barrett and Li, 2002).*

[4] *This model has been further extended in Spiller and Wood (1988), Sexton et al. (1991), Baulch (1997), Kleit (1998, 2001), Bailey (1998), Barrett and Li (2002) and Negassa and Myers (2007).*

parameters are held constant throughout the entire observation period. As these assumptions may look too restrictive in applications based on daily data, an enriched dynamic specification is also detailed in the paper.

We believe that this framework can provide useful guidance to a large audience interested in the functioning of the restructured natural gas industries (e.g., competition authorities, regulators, market analysts), and to researchers engaged in the detailed modeling of these industries.[5] As an application, we examine the spatial arbitrages performed between the two oldest European markets for wholesale natural gas in Belgium and the UK. This allows us to present a series of original empirical findings that: (i) show that all the arbitrage opportunities between the two zones are being exploited, but (ii) confirm the presence of market power in the spatial arbitrages. As the detailed institutional arrangements created for these two markets have largely shaped the designs of the other Continental markets, we believe that these findings provide a valuable contribution to the policy debate related to the restructuring of the European market for natural gas.

Despite the importance of market power concerns in the energy policy debates, the market power potentially exerted by natural gas arbitragers has hitherto been little studied. A notable exception is Rupérez Micola and Bunn (2007) who apply standard regression techniques to examine the relationship between the pipeline capacity utilization (i.e., the ratio of utilized to maximum capacity) and the absolute price difference between Belgium and the UK. Their results document the presence of market splitting at moderate levels of capacity utilization which, according to the authors, suggests the presence of market power inefficiencies. However, neither the direction of the trade flows nor the intermarket transfer costs play any role in their analysis. By taking these features into account, our paper confirms the presence of market power, even if all the arbitrage opportunities are being exploited, and connects the empirical results to the theoretical literature on spatial price determination.

The remaining sections of this paper are organized as follows. Section 2 details the theoretical conditions for spatial equilibrium between two markets linked by a capacity-constrained transportation infrastructure. Section 3 presents an adapted empirical methodology to investigate whether these conditions hold or not. Then, Section 4 details an application of this methodology to the case of the Interconnector UK, a natural gas pipeline connecting the UK to Continental Europe. Finally, the last section offers a summary and some concluding remarks.

2. Theoretical background

The empirical method described in the next section explicitly refers to the theoretical conditions for spatial equilibrium between two markets connected by a capacity-constrained transportation

[5] *In recent years, there has been an upsurge of interest in the application of operations research techniques to simulate the functioning of a restructured natural gas industry with the help of large scale equilibrium models (cf. the special issue of the Energy Journal: Huntington, 2009). However, the selection of the behavioral assumptions used to represent spatial arbitrages in these numerical models remains controversial as some models posit the existence of competitive spatial arbitrages (e.g. Golombek et al., 1995), whereas imperfect competition assumptions are used in others (e.g., Abada et al., 2013).*

infrastructure. This section presents these conditions, and introduces the notation used in the rest of the paper. We make two polar assumptions regarding the traders' aggregate behavior: perfect competition and monopolistic behavior. We present the short-run spatial equilibrium conditions in these two cases.

We consider two markets i and j located in different regions that trade a homogeneous commodity. We aim to analyze the direction-specific arbitrages that can be performed from market j to market i at time t. These arbitrages are based on a single transportation infrastructure that has a direction-specific, finite capacity K_{jit} that can change over time.[6] The local supply of that commodity is assumed to be competitive in both markets. We assume that, in each market, the aggregate total production cost is a convex, twice-continuously differentiable function. For each region i at time t, we assume that there is a linear inverse demand function: $p_{it}(q) = a_{it} - b_i q$ where a_{it} is the intercept and b_i is a strictly positive slope coefficient.[7] We ignore price uncertainty and respectively denote P_{it} and P_{jt} the local market clearing prices in each location.

The transportation infrastructure is owned by a regulated infrastructure company that sells transportation rights to a set of identical trading firms. The trading firms' unique activity is to perform spatial arbitrages. A transportation right provides its owner with the right to transfer up to one unit of good from market j to market i at each time period t during the infrastructure's lifetime. At any time t, the operating cost incurred by the infrastructure company is assumed to be recovered from the transportation rights owners in proportion to their use of the infrastructure at that time. Thus, the infrastructure company charges a non-discriminatory price per unit of good transported. The capital cost of this infrastructure is assumed to be recovered in lump sum charges paid by the transportation rights owners and we do not consider this cost further in the analysis. The total arbitrage cost, including all the transportation costs and the transaction costs incurred by a rights owner when performing an arbitrage from market j to market i at time t is assumed to be a linear function of the trade flow at that time. The associated marginal arbitrage cost is denoted τ_{jit}. We assume that there are no transport lags so that spatial arbitrage can take place within each observation period. The non-negative aggregate trade flow from j to i measured at time t is denoted Q_{jit}.

a – Case A: Perfectly competitive spatial arbitrages

In this case, we assume that traders adopt a price-taking behavior at each location. At time t, their aggregate behavior can be modeled using the following profit-maximization problem:

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[6] These variations are caused by exogenously determined fluctuations in a series of factors including: the operating pressures of the adjacent national pipeline systems, the flow temperature, the chemical composition of the natural gas. See Chenery (1949) and Massol (2011) for an engineering-based introduction to gas pipeline economics.

[7] We can remark that these slope coefficients are not subscripted with the time index and are thus assumed to be constant. In contrast, the intercepts of these inverse demand functions are assumed to be time-varying parameters (because of the seasonal variations observed in natural gas demand). These assumptions are frequently used in the context of restructured electricity markets (e.g., Day and Bunn, 2001).

$$\underset{Q_{jit}}{Max} \qquad \Pi^{C}_{jit}\left(Q_{jit}\right)=\left(P_{it}-P_{jt}-\tau_{jit}\right)Q_{jit} \tag{1}$$

$$\text{s.t.} \qquad Q_{jit} \le K_{jit} \tag{2}$$

$$Q_{jit} \ge 0 \tag{3}$$

where: the objective function (1) represents the total profits obtained by these trading firms, condition (2) describes the transportation capacity constraint, and condition (3) indicates that the trade flow from market j to market i must be non-negative.

Denoting ξ_{jit} the dual variable associated with the transportation capacity constraint (2), the Karush-Kuhn-Tucker conditions of this constrained optimization problem are:

$$0 \le Q_{jit}, \qquad P_{it}-P_{jt}-\tau_{jit}-\xi_{jit} \le 0 \quad \text{and} \quad \left(P_{it}-P_{jt}-\tau_{jit}-\xi_{jit}\right)Q_{jit}=0, \tag{4}$$

$$0 \le \xi_{jit}, \qquad Q_{jit} \le K_{jit} \quad \text{and} \quad \left(Q_{jit}-K_{jit}\right)\xi_{jit}=0. \tag{5}$$

These two complementarity conditions together define the equilibrium conditions for competitive spatial arbitrages at time t.

The marginal profit to spatial arbitrages is equal to the difference between the market-clearing price at location i and the sum of the price at location j and τ_{jit} the marginal arbitrage cost related to trade. The dual variable ξ_{jit} can be interpreted as a marginal congestion cost. The complementarity condition (5) ensures that the marginal congestion cost ξ_{jit} is equal to zero whenever the transportation capacity constraint (2) is slack, and that ξ_{jit} is positive when this constraint is binding. In case of a zero marginal congestion cost (i.e., $\xi_{jit}=0$), the complementarity condition (4) is fully consistent with the logic of the Enke-Samuelson-Takayama-Judge spatial equilibrium models because it ensures: (i) that there is no trade from market j to market i (i.e., $Q_{jit}=0$) when the marginal profit to spatial arbitrage is negative, and (ii) that the marginal profit to spatial arbitrage is zero when trade occurs and it is not constrained by the infrastructure's capacity (i.e., $0<Q_{jit}<K_{jit}$). In our setup, we allow for a binding capacity constraint (i.e., $Q_{jit}=K_{jit}$) which, according to the complementary condition (4), ensures that the marginal profit to spatial arbitrage is positive (i.e., $P_{it}-P_{jt}-\tau_{jit}\ge 0$). In this case, there exists a scarcity rent $\left(P_{it}-P_{jt}-\tau_{jit}\right)K_{jit}$ that accrues to the traders.

<u>b – Case B: Monopolistic spatial arbitrages</u>

We now assume, at the other extreme, that traders collectively behave as a monopoly, i.e., that they know how the prices in each region react to the quantities supplied. At time t, their aggregate behavior can be modeled using the following profit-maximization problem:

$$\underset{Q_{jit}}{Max} \quad \Pi_{jit}^{M}\left(Q_{jit}\right)=\left(p_{it}\left(S_{it}+Q_{jit}\right)-p_{jt}\left(S_{jt}-Q_{jit}\right)-\tau_{jit}\right)Q_{jit} \tag{6}$$

$$\text{s.t.} \quad Q_{jit} \leq K_{jit} \tag{7}$$

$$Q_{jit} \geq 0 \tag{8}$$

where: $p_{it}(.)$ and $p_{jt}(.)$ are the local inverse demand functions, and S_{it} and S_{jt} are the aggregate supplies decided by all the local producers at each location.

We denote again ξ_{jit} the Lagrange multiplier associated with the transportation capacity constraint (7), and P_{it} and P_{jt} the local market clearing prices. The Karush-Kuhn-Tucker conditions of this constrained optimization problem are:

$$0 \leq Q_{jit}, \quad P_{it}-P_{jt}-\tau_{jit}-\left(b_i+b_j\right)Q_{jit}-\xi_{jit} \leq 0 \text{ and } \left(P_{it}-P_{jt}-\tau_{jit}-\left(b_i+b_j\right)Q_{jit}-\xi_{jit}\right)Q_{jit}=0 \tag{9}$$

$$0 \leq \xi_{jit}, \quad Q_{jit} \leq K_{jit} \quad \text{and} \quad \left(Q_{jit}-K_{jit}\right)\xi_{jit}=0 \tag{10}$$

These two complementarity conditions together define the equilibrium conditions for monopolistic spatial arbitrages at time t.

The economic interpretation of the complementarity conditions (9) and (10) is similar to those detailed for the case of competitive arbitrages except the marginal profit to spatial arbitrage is now equal to $P_{it}-P_{jt}-\tau_{jit}-\left(b_i+b_j\right)Q_{jit}$. Thus, in the event of monopolistic arbitrages, the spatial price differential is always larger than the marginal arbitrage cost when trade is observed (even when the congestion constraint is slack). This reflects the traders' ability to exert market power by restricting intermarket trade to generate some monopoly rents.

3. Methodology

This section presents the methodology used in this manuscript. We first adapt the existing static PBM framework to take into account the role of both pipeline capacity constraints and market power. Subsequently, we detail the empirical specification and a dynamic extension to the static model.

3.1 An adapted parity bounds model

We now define seven mutually exclusive trade regimes and relate them to the theoretical conditions for spatial equilibrium detailed in the previous section. In addition to the six trade regimes considered in the PBM proposed in Barrett and Li (2002), we introduce a new one that takes into account the case of pipeline congestion. Moreover, for each of these trade regimes, we distinguish between the two polar cases of perfectly competitive and monopolistic spatial arbitrages.

As shown in Table 1, marginal profits to spatial arbitrage and trade flow considerations can be combined to define a taxonomy of trade regimes governing the arbitrages from market j to market i.

Regarding marginal profits to spatial arbitrage, three basic states can be defined depending on the value of these marginal profits: zero, strictly positive, and strictly negative. Regarding trade flows, two basic states can be identified depending on whether a positive trade flow is observed or not. Following Barrett and Li (2002), each of these six regimes is labeled I to VI, where odd numbers are used for regimes with strictly positive trade flows and even numbers for those without trade.

Table 1. The trade regimes in each direction

	Trade is observed: $0 < Q_{jit} \leq K_{jit}$	No trade is observed: $Q_{jit} = 0$
zero marginal profits to spatial arbitrage	Regime I λ_{I}	Regime II λ_{II}
positive marginal profits to spatial arbitrage	Regime III_a iff $Q_{jit} < K_{jit}$ λ_{III_a} Regime III_b iff $Q_{jit} = K_{jit}$ λ_{III_b}	Regime IV λ_{IV}
negative marginal profits to spatial arbitrage	Regime V λ_{V}	Regime VI λ_{VI}

In regimes I and II, the marginal profit to spatial arbitrage is equal to 0. As shown in the previous section, depending on the assumption posited for the behavior of the trading sector, one of the following conditions is binding:

Case A: Competitive arbitrages

$$P_{it} - P_{jt} - \tau_{jit} = 0 \tag{11}$$

Case B: Monopolistic arbitrages

$$P_{it} - P_{jt} - \tau_{jit} - (b_i + b_j) Q_{jit} = 0 \tag{12}$$

In case of price-taking behavior (Case A), the spatial price differential is equal to the marginal transfer cost. In case of monopolistic arbitrages (Case B), the possibility to exert market power results in a spatial price differential that exceeds the marginal transfer cost and the difference between the two is proportional to the observed trade flow. In Case A (respectively B), each of the two regimes verifies the complementarity slackness condition (4) (respectively (9)) when there is no congestion cost (i.e., $\xi_{jit} = 0$). Therefore, both regimes are consistent with the conditions for a spatial equilibrium.

In regimes III and IV, the marginal profit to arbitrage from j to i is strictly positive:

Case A: Competitive arbitrages

$$P_{it} - P_{jt} - \tau_{jit} > 0 \tag{13}$$

Case B: Monopolistic arbitrages

$$P_{it} - P_{jt} - \tau_{jit} - (b_i + b_j) Q_{jit} > 0 \tag{14}$$

In both of these regimes, markets are separated and there are unseized opportunities for profitable spatial arbitrage. Still, in case of positive trade (regime III), the observed insufficient arbitrages might result from the capacity-constrained nature of the transportation infrastructure. Indeed, the complementarity conditions detailed in the preceding section indicate that, in case of a binding capacity constraint (i.e., $Q_{jit} = K_{jit}$), observing a strictly positive value for the marginal profit to arbitrage is consistent with the conditions for a short-run spatial equilibrium. In contrast, the joint observation of strictly positive marginal profit to arbitrages and a slackening in the infrastructure's capacity constraint violate the conditions for a spatial equilibrium. Thus, we propose a modification to the original model and further decompose regime III into two mutually exclusive regimes labeled III_a and III_b. In regime III_a, the observed trade flows verify $0 < Q_{jit} < K_{jit}$ whereas a binding capacity constraint (i.e., $Q_{jit} = K_{jit}$) is observed in regime III_b. Therefore, the latter regime, but not the former, is consistent with the conditions for a spatial equilibrium.

In regimes V and VI, the marginal profit to arbitrage from j to i is strictly negative:

Case A: Competitive arbitrages

$$P_{it} - P_{jt} - \tau_{jit} < 0 \tag{15}$$

Case B: Monopolistic arbitrages

$$P_{it} - P_{jt} - \tau_{jit} - \left(b_i + b_j\right) Q_{jit} < 0 \tag{16}$$

In both regimes, there are no profitable arbitrage opportunities. In regime VI, trade is not occurring and the observed local prices correspond to autarky prices. This regime is consistent with the conditions for a spatial equilibrium. In contrast, regime V indicates that trade is occurring despite negative marginal profits which are not consistent with equilibrium conditions.

In sum, having introduced a further distinction between regimes III_a and III_b, a total of seven regimes are thus considered in our analysis. The estimated probability to observe regime r is denoted λ_r. Spatial equilibrium conditions hold with probability $\left(\lambda_I + \lambda_{II} + \lambda_{III_b} + \lambda_{VI}\right)$ and the estimated probability to observe disequilibrium is $\left(\lambda_{III_a} + \lambda_{IV} + \lambda_V\right)$.

3.2 Empirical specification

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, the observable marginal arbitrage cost, the trade flow, and the available transportation capacity.

From an empirical perspective, the marginal arbitrage cost τ_{jit} can be decomposed into two components: an observable portion named T_{jit} (i.e., a series that comprises all the transportation and transaction costs available to the modeler), and an unobservable one which is assumed to be explained by a constant and by a vector of observable exogenous factors Z_{jit}.[8] So, in what follows, the marginal

[8] *Consistent with our assumption of total arbitrage costs that vary linearly with the trade flows, this vector does not include the trade flow variable.*

arbitrage cost is assumed to be $\tau_{jit} \equiv T_{jit} + \alpha_{ji} + Z_{jit}\beta_{ji}$, where α_{ji} and β_{ji} are direction-specific parameters to be estimated.[9]

In case of monopolistic arbitrages, the sum of the slope coefficients for the local inverse demand functions is unlikely to be readily available to the modeler. So, we introduce γ an unknown parameter to be estimated that will be interpreted as $\left(b_i + b_j\right)$ the sum of the slope coefficients. So, we expect the estimated value for γ to be non-negative.

Denoting $R_{jit} \equiv P_{it} - P_{jt} - T_{jit}$ the series that represents the observable portion of the marginal rent to spatial arbitrage, the marginal profits to arbitrage in each of the three distinct cases (zero, positive and negative) are modeled using the following switching regression model (Sexton et al., 1991; Baulch, 1997; Barrett and Li, 2002):

<u>Case A: Competitive arbitrages</u>

Regimes I & II:

$$R_{jit} - \left(\alpha_{ji} + Z_{jit}\beta_{ji}\right) = \varepsilon_{jit} \tag{17}$$

Regimes III_a, III_b & IV:

$$R_{jit} - \left(\alpha_{ji} + Z_{jit}\beta_{ji}\right) = \varepsilon_{jit} + \mu_{jit} \tag{18}$$

Regimes V & VI:

$$R_{jit} - \left(\alpha_{ji} + Z_{jit}\beta_{ji}\right) = \varepsilon_{jit} - \upsilon_{jit} \tag{19}$$

<u>Case B: Monopolistic arbitrages</u>

Regimes I & II:

$$R_{jit} - \left(\alpha_{ji} + Z_{jit}\beta_{ji}\right) - Q_{jit}\gamma = \varepsilon_{jit} \tag{20}$$

Regimes III_a, III_b & IV:

$$R_{jit} - \left(\alpha_{ji} + Z_{jit}\beta_{ji}\right) - Q_{jit}\gamma = \varepsilon_{jit} + \mu_{jit} \tag{21}$$

Regimes V & VI:

$$R_{jit} - \left(\alpha_{ji} + Z_{jit}\beta_{ji}\right) - Q_{jit}\gamma = \varepsilon_{jit} - \upsilon_{jit} \tag{22}$$

where: α_{ji}, β_{ji} and γ are unbounded real-valued parameters; ε_{jit} is a random error that is assumed to be i.i.d. normally distributed with a zero mean and variance σ_ε^2; and μ_{jit} and υ_{jit} are i.i.d. random samples from zero-centered normal distributions truncated above at 0 with respective variance parameters σ_μ^2 and σ_υ^2.

The specifications used to model the cases of competitive and monopolistic spatial arbitrages differ only in the markup term $Q_{jit}\gamma$. Thus, a statistical test of the null hypothesis $\gamma = 0$ (e.g., a likelihood ratio test) can be conducted to test the null hypothesis of perfectly competitive spatial arbitrages. For the sake of brevity, only the unrestricted model based on equations (20), (21), and (22) is detailed hereafter.

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[9] *From an empirical perspective, this interpretation fails to acknowledge data quality issues in the time series used in the estimation procedure. As argued in Barrett and Li (2002), the constant* α_{ji} *can also reflect any measurement bias in either the price or the observable transaction cost series. We proceed with this caveat in mind, assuming the validity of this interpretation.*

Denoting $\theta \equiv (\alpha_{ji}, \beta_{ji}, \gamma, \sigma_{\varepsilon}, \sigma_{\mu}, \sigma_{\upsilon})$ the parameter vector to be estimated and $\pi_{jit} \equiv R_{jit} - (\alpha_{ji} + Z_{jit}\beta_{ji}) - Q_{jit}\gamma$ the random variable that gives the marginal profit from spatial arbitrage at time t, the joint density function for the observation at time t is the mixture distribution:

$$\begin{aligned} f_{jit}\left(\pi_{jit} \middle| (\lambda,\theta)\right) \equiv\ & A_{jit}\left[\lambda_I f_{jit}^{I}\left(\pi_{jit}|\theta\right) + \left((1-B_{jit})\lambda_{III_a} + B_{jit}\lambda_{III_b}\right) f_{jit}^{III}\left(\pi_{jit}|\theta\right) + \lambda_V f_{jit}^{V}\left(\pi_{jit}|\theta\right)\right] \\ & + (1-A_{jit})\left[\lambda_{II} f_{jit}^{II}\left(\pi_{jit}|\theta\right) + \lambda_{IV} f_{jit}^{IV}\left(\pi_{jit}|\theta\right) + \lambda_{VI} f_{jit}^{VI}\left(\pi_{jit}|\theta\right)\right] \end{aligned} \tag{23}$$

where: A_{jit} is an indicator variable that takes a value of 1 if trade is observed and zero otherwise; B_{jit} is an indicator variable that takes a value of 1 if the transportation infrastructure is congested and zero otherwise; $f_{jit}^{I}(\pi_{jit}|\theta)$ and $f_{jit}^{II}(\pi_{jit}|\theta)$ are normal density functions; $f_{jit}^{III}(\pi_{jit}|\theta)$ and $f_{jit}^{IV}(\pi_{jit}|\theta)$ (respectively $f_{jit}^{V}(\pi_{jit}|\theta)$ and $f_{jit}^{VI}(\pi_{jit}|\theta)$) are the density functions derived in Weinstein (1964) for the sum of a normal random variable and a centered-normal random variable truncated above (respectively below) at 0.

Denoting λ the vector of the estimated probabilities to observe the seven regimes, the likelihood function for a sample of observations $\{R_{jit}, Z_{jit}, Q_{jit}, K_{jit}\}$ is:

$$L(\lambda,\theta) \equiv \prod_{t=1}^{N} f_{jit}\left(\pi_{jit} \middle| (\lambda,\theta)\right) \tag{24}$$

The model can be estimated by maximizing the logarithm of the likelihood function with respect to regime probabilities and model parameters subject to the constraints that the regime probabilities sum to one and that each of these probabilities lies in the unit interval.

This specification differs from that of Barrett and Li (2002) in four ways. First, we show how a parity bound model can be used to test the null hypothesis of competitive spatial arbitrages. Second, contrary to Barrett and Li, the two markets under scrutiny are connected by a capacity-constrained transportation infrastructure. So, a seventh regime labeled III_b is introduced to account for the explanatory role played by infrastructure congestion issues in the observation of positive marginal profits to spatial arbitrage. Third, we introduce a vector of observable exogenous factors Z_{jit} to capture the effects of omitted arbitrage costs. This specification provides fuller information to analyze market relationships. Lastly, we follow Baulch (1997) and Negassa and Myers (2007) and relax the restriction to use a unique variance parameter for the half-normal distributions of the two non-negative error terms μ_{ijt} and υ_{ijt} in the likelihood functions.[10]

[10] *One might argue that, in certain applications, sample size considerations could motivate the use of the restriction* $\sigma_{\mu} = \sigma_{\upsilon}$ *to reduce the number of parameters to be estimated. Yet, this restriction can hardly be justified a priori.*

3.3 Dynamic extension

The specification above has a static nature. So, we now detail an extended version aimed at capturing the inter-period linkages that may be observed in commodity markets.

a – 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).[11] As the presence of unmodeled autocorrelation can result in inefficient estimates, the presence of serial correlation has to be appropriately corrected for.[12]

In this paper, we apply a Bayesian approach that is similar to the one in Kleit (2001). We aim to extend the Barrett and Li framework to adjust for the presence of serial correlation in the error term ε_{jit}. Yet, a difficulty emerges: the exact value $\varepsilon_{ji(t-1)}$ cannot be directly observed. However, we can consider the expected value of $\varepsilon_{ji(t-1)}$, given the evidence provided by the previous observation, which results in the modified specification:

Regimes I & II: $$R_{jit} - \left(\alpha_{ji} + Z_{jit}\beta_{ji}\right) - Q_{jit}\gamma - \rho_{ji}\mathrm{E}\left(\varepsilon_{ji(t-1)} \middle| \eta_{ji(t-1)}\right) = \varepsilon_{jit} \quad (25)$$

Regimes $\mathrm{III_a}$, $\mathrm{III_b}$ & IV: $$R_{jit} - \left(\alpha_{ji} + Z_{jit}\beta_{ji}\right) - Q_{jit}\gamma - \rho_{ji}\mathrm{E}\left(\varepsilon_{ji(t-1)} \middle| \eta_{ji(t-1)}\right) = \varepsilon_{jit} + \mu_{jit} \quad (26)$$

Regimes V & VI: $$R_{jit} - \left(\alpha_{ji} + Z_{jit}\beta_{ji}\right) - Q_{jit}\gamma - \rho_{ji}\mathrm{E}\left(\varepsilon_{ji(t-1)} \middle| \eta_{ji(t-1)}\right) = \varepsilon_{jit} - \upsilon_{jit} \quad (27)$$

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

We now show how the expected value $\mathrm{E}\left(\varepsilon_{ji(t-1)} \middle| \eta_{ji(t-1)}\right)$ can be computed. Given the observed value of the lagged residual $\eta_{ji(t-1)}$ and the parameter vector $\theta_1 \equiv \left(\theta, \rho_{ji}\right)$, Bayes theorem can be

[11] *In the application discussed below (an analysis of market integration between two markets connected by a pipeline system), at least two engineering-based arguments can motivate the presence of autocorrelation. First, from a dynamic perspective, a pipeline system can be described as a slow-moving transportation infrastructure as a couple of hours are typically needed to move a given molecule of methane from one market to the other. Second, the operation of a natural gas pipeline system can create temporary energy storage (the so-called line-pack buffer). As a result, daily observations are likely to jointly represent the outcome of decisions taken both today and yesterday.*

[12] *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.*

invoked to evaluate $\mathrm{P}_{t-1}^{r} \equiv \mathrm{P}_{t-1}\left(r \middle| \eta_{ji(t-1)}, \theta_1\right)$ the posterior probability that the residual observed at time $t-1$ was generated by regime r (Kiefer, 1980; Spiller and Wood, 1988):

$$\mathrm{P}_{t-1}^{r} = \frac{\lambda_r f_{ji(t-1)}^{r}\left(\eta_{ji(t-1)} \middle| \theta_1\right)}{\left(\lambda_{III_a} + \lambda_{III_b}\right) f_{ji(t-1)}^{III}\left(\eta_{ji(t-1)} \middle| \theta_1\right) + \sum_{\substack{k=I \\ k \neq III}}^{VI} \lambda_k f_{ji(t-1)}^{k}\left(\eta_{ji(t-1)} \middle| \theta_1\right)} . \tag{28}$$

The expected value $\mathrm{E}\left(\varepsilon_{ji(t-1)} \middle| \eta_{ji(t-1)}\right)$ can be constructed from the observed residual $\eta_{ji(t-1)}$ by: (i) subtracting $\mathrm{E}(\mu)$ the expected value of the one-sided random variable μ_{jit} weighted by the posterior probability to observe the regimes $\mathrm{III_a}$, $\mathrm{III_b}$ or IV; and, (ii) adding $\mathrm{E}(\upsilon)$ the expected value of the non-negative half-normal random variable υ_{jit} weighted by the posterior probability to observe the regimes V or VI,[13] that is:

$$\mathrm{E}\left(\varepsilon_{ji(t-1)} \middle| \eta_{ji(t-1)}\right) = \eta_{ji(t-1)} - \left[\mathrm{P}_{t-1}^{III_a} + \mathrm{P}_{t-1}^{III_b} + \mathrm{P}_{t-1}^{IV}\right]\mathrm{E}(\mu) + \left[\mathrm{P}_{t-1}^{V} + \mathrm{P}_{t-1}^{VI}\right]\mathrm{E}(\upsilon) . \tag{29}$$

The construction of $\mathrm{E}\left(\varepsilon_{ji(t-1)} \middle| \eta_{ji(t-1)}\right)$ 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 $-1 < \rho_{ji} < 1$.[14]

<u>b – An adapted GARCH specification</u>

Regimes I and II model the cases of zero marginal profit to spatial arbitrage. In these regimes, the random variable representing the marginal profit to spatial arbitrage is assumed to be equal to the stochastic error term ε_{jit} which has the same finite variance σ_ε^2 for all observations. Yet, one may question the relevance of this homoscedastic assumption. A large empirical literature has documented the tendency of commodity prices to exhibit time-varying volatilities. Accordingly, the spatial price differential (and thus the marginal profit to spatial arbitrage) is likely to show signs of heteroscedasticity. Therefore, we now detail a modified specification whereby the variance of the marginal profit to spatial arbitrage observed in regimes I and II is allowed to vary over time.

For the purpose of capturing the dynamics of uncertainty, a Generalized Autoregressive Conditional Heteroscedasticity (GARCH) model (Bollerslev, 1986) represents an attractive approach

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[13] Denoting ϕ the density function of the standard normal distribution and Φ its cumulative distribution function, these expected values are: $\mathrm{E}(\mu) \equiv \sigma_\mu \phi(0) / (1 - \Phi(0))$ and $\mathrm{E}(\upsilon) \equiv \sigma_\upsilon \phi(0) / (1 - \Phi(0))$.

[14] Regarding the particular case of the first observation, an arbitrary value has to be taken for $\mathrm{E}\left(\varepsilon_{ji0} \middle| \eta_{ji0}\right)$ because η_{ji0} cannot be observed. In this paper, the initial value $\mathrm{E}\left(\varepsilon_{ji0} \middle| \eta_{ji0}\right)$ is taken as equal to zero (that is, the conditional mean of ε_{jit} given $\varepsilon_{ji(t-1)}$).

that has been widely applied to model commodity markets. Given the time series R_{jit}, Z_{jit} and Q_{jit} defined above, a GARCH(1,1) specification can be written as follows:

Regimes I & II: $$R_{jit} - \left(\alpha_{ji} + Z_{jit}\beta_{ji}\right) - Q_{jit}\gamma - \rho_{ji}\mathrm{E}\left(\varepsilon_{ji(t-1)} \middle| \eta_{ji(t-1)}\right) = \varepsilon_{jit} \tag{30}$$

Regimes $\mathrm{III_a}$, $\mathrm{III_b}$ & IV: $$R_{jit} - \left(\alpha_{ji} + Z_{jit}\beta_{ji}\right) - Q_{jit}\gamma - \rho_{ji}\mathrm{E}\left(\varepsilon_{ji(t-1)} \middle| \eta_{ji(t-1)}\right) = \varepsilon_{jit} + \mu_{jit} \tag{31}$$

Regimes V & VI: $$R_{jit} - \left(\alpha_{ji} + Z_{jit}\beta_{ji}\right) - Q_{jit}\gamma - \rho_{ji}\mathrm{E}\left(\varepsilon_{ji(t-1)} \middle| \eta_{ji(t-1)}\right) = \varepsilon_{jit} - \upsilon_{jit} \tag{32}$$

$$\varepsilon_{jit} = h_{jit} e_{jit} \tag{33}$$

$$h_{jit}^2 = \varpi_{ji} + \delta_{ji}\left[\mathrm{E}\left(\varepsilon_{ji(t-1)} \middle| \eta_{ji(t-1)}\right)\right]^2 + \varphi_{ji} h_{ji(t-1)}^2 \tag{34}$$

where: (30), (31) and (32) are the mean equations; (34) is the conditional variance equation; (33) relates the random error ε_{jit} to the standardized residual e_{jit} which is assumed to be an i.i.d. standard normal random variable; and ϖ_{ji}, δ_{ji} and φ_{ji} are the usual, non-negative, GARCH(1,1) parameters.

Compared to the usual GARCH specification, equation (34) involves the use of the squared expected value $\left[\mathrm{E}\left(\varepsilon_{ji(t-1)} \middle| \eta_{ji(t-1)}\right)\right]^2$ in spite of the true value $\varepsilon_{ji(t-1)}^2$ which cannot be observed in this regime switching model. Again, the construction of the expected value $\mathrm{E}\left(\varepsilon_{ji(t-1)} \middle| \eta_{ji(t-1)}\right)$ is based on a Bayesian approach. Given the observed value $\eta_{ji(t-1)}$, the values of the parameters $\theta_2 \equiv \left(\theta_1, \varpi_{ji}, \delta_{ji}, \varphi_{ji}\right)$ and $h_{ji(t-2)}^2$, we can evaluate the posterior probabilities P_{t-1}^r and thus $\mathrm{E}\left(\varepsilon_{ji(t-1)} \middle| \eta_{ji(t-1)}\right)$ using (29).

4. Application

4.1 Background

This application focuses on the so-called Interconnector (hereafter abbreviated to IUK), a bi-directional natural gas pipeline system connecting the UK National Transportation System (using the Bacton Terminal) to Zeebrugge (Belgium). This infrastructure allows spatial arbitrages between Europe’s two oldest spot markets for natural gas: (i) the UK’s NBP, which allows counterparties to trade a standardized lot of natural gas piped via the UK National Transmission System with a delivery point at the so-called National Balancing Point (NBP); and (ii) the Zeebrugge local market in Belgium, which is labeled ZEE.

We consider the period covering October 1, 2003, to October 5, 2006.[15] During that period, the IUK pipeline was the unique infrastructure linking the UK and Continental natural gas markets. In addition, that period corresponds to a steady institutional environment with unchanged access rules for both the IUK and the adjacent national pipeline systems. These features make the IUK case an attractive experiment to investigate the efficiency of the spatial price arbitrages that can be performed in a deregulated natural gas industry.

4.2 Data

We use daily transaction price data for day-ahead wholesale natural gas traded during working days as published by Platt's, a price-reporting service. For each working day (i.e., Monday to Friday), they reflect the price range of a standardized quantity of natural gas to be delivered at a constant flow rate throughout the next working day after assessment (e.g., Friday's assessment reflects Monday's delivery).[16] All prices are denominated in €/MWh. Given the extremely limited liquidity of within-day markets, we follow the usual convention and refer to these day-ahead prices as "spot" since they provide traders with a final opportunity to trade gas out of a forward position before physical delivery.

The fuel used by the IUK operator to power its compressor equipment represents the observable portion of the marginal arbitrage costs incurred by the traders. This is a direction-specific cost since, according to the pipeline operator, fuel gas consumption amounts to 0.8% of the quantity of gas transported when natural gas is piped from the UK to Belgium, and to 0.26% of the quantity of gas transported in the other direction. This fuel cost is evaluated using the price of natural gas in the exporting market.

The unobservable portion of the marginal arbitrage costs includes two kinds of costs: (i) the entry (respectively, exit) charges incurred by the traders who import gas to (respectively export gas from) the UK, and (ii) all the unobserved transaction costs (e.g., unmeasured transactions costs, information gaps...). Hereafter, these costs will be estimated using a constant, a time trend, and two dummy variables: $D_{2004-2005}$ that takes the value 1 during the period covering October 1, 2004, to September 30, 2005, and $D_{2005-2006}$ that takes the value 1 after October 1, 2005. Each period corresponds to a "standard gas year" during which the regulated Entry-Exit tariff system used by the UK National Transportation System is kept unchanged.

Regarding trade flow data, the wish may be to use an aggregate variable gathering all the transportation nominations communicated at the end of any working day for delivery during the next working day. Unfortunately, these data are confidential. So, this study uses a proxy: a historical flow series representing the physical daily flow of natural gas, measured in GWh/day, that transited through

[15] *This starting date has been chosen to omit the number of partial closures that happened during the summer of 2003. Moreover, the transited flows of natural gas were almost unidirectional (from the UK to the Continent) before that date (Futyan, 2006). This terminal date corresponds to the opening of the Langeled infrastructure, a pipeline system that together with already existing offshore pipelines, allowed Norwegian gas producers to perform spatial arbitrages between the UK and the Continent, thereby offering an alternative to the IUK.*

[16] *Further information on the methodology used to construct these market-on-close price assessments is available in Platt's (2012).*

the IUK as reported on the pipeline operator's website.[17] Thus, we proceed under the assumption that the physical gas flow measured during a given working day represents an unbiased estimator of the aggregate transportation nominations decided during the previous working day (at the time when trade occurs in the corresponding day-ahead market).[18]

According to the Interconnector operator, the nominal transportation capacity from the UK to Belgium remained unchanged during the entire sample period. In the other direction, the installation of some compressor equipment in Zeebrugge on November 8, 2005, increased the transportation capacity. Unfortunately, information related to the available daily transportation capacities remains unavailable. So, we follow Rupérez Micola and Bunn (2007) and consider the historical maximum values of the trade flows. The historical maxima were: 624.63 GWh/day from the UK to Belgium, and 310.24 GWh/d prior to November 8, 2005, (respectively 511.80 GWh/d after that date) in the other direction.

The daily flow capacity of a point-to-point natural gas pipeline is a time-varying parameter that depends on a series of exogenous factors (e.g., the operating pressures of the adjacent national pipeline systems, the flow temperature, the chemical composition of the natural gas). Hence, the historical maximum daily flow cannot necessarily be attained. We proceed by assuming that congestion is likely to be a source of concern when the observed capacity utilization ratio (measured against the historically maximum) exceeds 80%. Hereafter, this threshold is used to distinguish regimes III_a and III_b.

The data set has been modified in two ways. First, both the Belgium and UK markets are closed on national bank holidays. To account for differences in the national calendars, all the observations related to a bank holiday in either Belgium or the UK have been disregarded in the subsequent analyses. Second, we excluded observations made on dates during which the Interconnector service was unavailable due to planned maintenance (these dates are documented on the IUK's website). As a result, we assembled time series data containing 723 daily observations on prices, compressor fuel costs, and trade flows in each direction.

4.3 Descriptive statistics

An examination of trade data indicates that out of these 723 observations, 369 correspond to net positive exports to Belgium (of which 26 correspond to a congested infrastructure), 341 to net imports to the UK (of which 46 correspond to a congested infrastructure) and 13 to zero trade.

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[17] *Cf. www.interconnector.com*

[18] *On a given working day, pipeline users are offered the possibility to revise the transportation service requested at the end of the previous working day. This is the so-called within-day re-nominations. Yet, for the pipeline operator, these within-day re-nominations generate a significant extra operation costs. As a result, the detailed pricing rules adopted by the pipeline operator have been explicitly designed to render these within-day re-nominations extremely costly. So, users have a strong incentive to contract their real transportation needs for day d+1 at the end of day d (i.e., before the close of the day ahead market). Therefore, we proceed assuming that these within-day re-nominations can be neglected.*

Table 2 details the correlation coefficients between the two local price series in both levels and first differences. Following Stigler and Sherwin (1985), these high values could be interpreted as positive signs of market integration, though the “degree” of that integration seems to be weaker when the natural gas is exported from the UK. So, a direction-specific approach seems to be needed to further investigate the degree of spatial integration between these two markets.

Table 2. Correlation coefficients for the two price series

	Entire sample	Subsample defined by net positive exports to Belgium	Subsample defined by net positive exports to the UK
correlation in levels	0.989***	0.985***	0.988***
correlation in first differences	0.900***	0.760***	0.905***

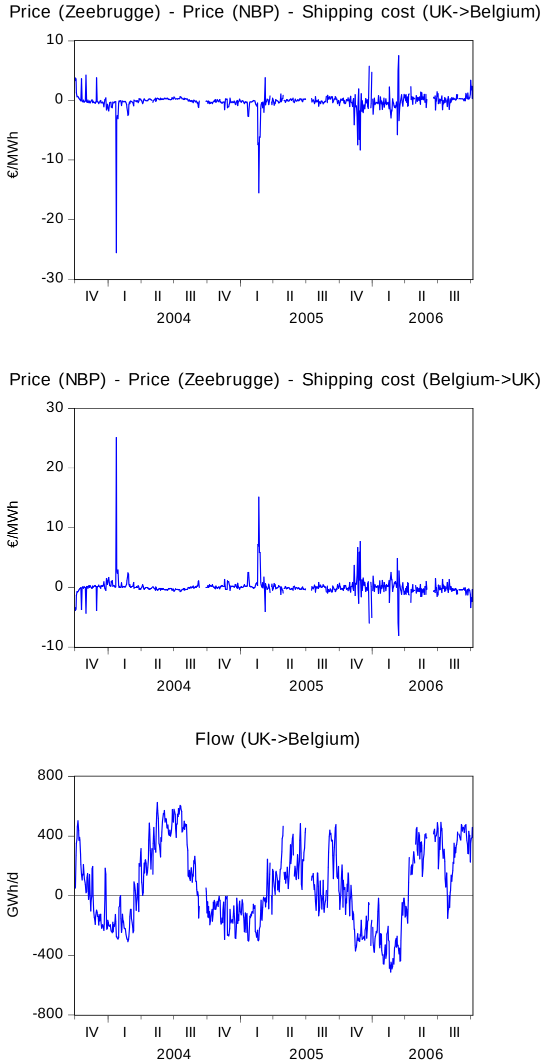
Note: *** indicate significance at the 0.01 level.

Table 3 summarizes the descriptive statistics for the two series R_{jit} (i.e., the observable portion of the marginal rent to spatial arbitrage). The distributional properties of these series show some signs of non-normality as a very large leptokurtosis is observed in both cases. Figure 1 provides plots of these two series and the measured pipeline flow from Bacton (UK) to Zeebrugge (Belgium). A visual inspection of these plots suggests the presence of volatility clustering which motivates the use of a specification that incorporates some GARCH features. The estimated first-order autocorrelation coefficients reveal clear evidence of serial correlation (0.311^{***} for the series $R_{NBP \to ZEE,t}$ and 0.301^{***} for the series $R_{ZEE \to NBP,t}$). This finding is in favor of a dynamic specification able to correct for serial correlation.

Table 3. Descriptive statistics for the marginal rent to spatial arbitrage

	Entire sample	
	$R_{NBP \to ZEE}$	$R_{ZEE \to NBP}$
Mean	-0.232	0.044
Median	-0.100	-0.051
Maximum	7.484	25.189
Minimum	-25.543	-8.096
Std. Dev.	1.581	1.557
Skewness	-7.394	7.396
Kurtosis	108.946	112.664
Jarque-Bera *(p-value)*	344730.200 *(0.000)*	368877.600 *(0.000)*
Observations	723	723

Figure 1. Data plots



4.4 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.[19]

To obtain a feasible starting point, we first consider the simplest possible static specification (i.e., omitting the exogenous explanatory variables Z_{jit} and assuming zero values for both the autocorrelation and the GARCH parameters). The converged solution for this restricted specification is then used as a feasible starting point for the unrestricted model. The optimization problem at hand has the potential for local maxima, which is a source of concern because the outcome of a non-linear programming solver may depend on the location of the starting point. To address this problem, the

[19] *All the estimates reported in this paper have been obtained using 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, and then a switch back to DFP for 20 iterations, and so forth.*

first solution is systematically compared to the ones obtained with a sample of 1,000 starting points uniformly drawn over a range of possible starting values. The converged solution that provides the highest likelihood value is systematically stored.

b – Empirical results

Estimation results are reported in Table 4. That table details the estimates obtained for: the unobserved marginal transaction costs (α, β_{time}, $\beta_{D_{2004-2005}}$, $\beta_{D_{2005-2006}}$), the market power coefficient (γ), the autocorrelation parameter (ρ), the regime probabilities (λ's), the GARCH parameters used to model the heteroscedasticity related to regimes I and II (ϖ, δ, φ), the standard deviation parameters for the truncated normal distributions (σ_μ, σ_υ), and a series of likelihood ratio tests.

From these estimation results, several facts stand out. First, the estimated values for the coefficient γ are positive, as expected. These estimates are highly significant in both directions, which reveals the presence of imperfectly competitive arbitrages across the Channel. A further confirmation is provided by the likelihood ratio tests: the null hypothesis of competitive arbitrages is firmly rejected in both directions. So, we cannot reject the assumption of imperfectly competitive arbitrages during that period. This finding is consistent with the results in Rupérez-Micola and Bunn (2007).

Second, the high estimates obtained for λ_I and λ_{II} in both directions reveal that the observed spatial price difference is predominantly explained by the sum of the unobserved marginal transaction costs and the markup term. These very high values result in a very high probability of observing an imperfectly competitive spatial market equilibrium. Following Barrett and Li (2002) the probability of spatial market equilibrium conditions holding is in the range defined by the minimum and the maximum values of the direction-specific sums $\left(\lambda_I + \lambda_{II} + \lambda_{III_b} + \lambda_{VI}\right)$, that is (0.9468, 0.9477).

Third, the probabilities λ_{III_b} to jointly observe infrastructure congestion and strictly positive marginal profits to spatial arbitrage regime are either zero or extremely low. These estimated values are consistent with the analysts' consensus summarized in Futyan (2006) on: (i) the oversized nature of the IUK's transportation capacity when natural gas is flowing to the Continent and (ii) the likely capacity-constrained nature of the IUK in the opposite direction (before the November 2005 capacity increase). In contrast, the estimated probabilities λ_{III_a} are larger and highly significant. Infrastructure congestion issues that are directly related to the Interconnector pipeline cannot be invoked to explain the presence of these strictly positive marginal profits to spatial arbitrage. These observed trade barriers could, for example, be due to pipeline congestion in the adjacent systems. Still, in relative terms, the probability of observing this regime is small.

Fourth, the marginal cost parameters included in the mean equation are significant in both directions, which confirms the presence of unobserved marginal transaction costs. The estimated unobserved marginal costs are strictly negative for all the observations, which suggests the presence of unobserved marginal benefits to trade.

Table 4. Estimation results for natural gas trade across the Channel

	From UK to Belgium	From Belgium to UK
Mean parameters		
α	-0.3164***	-0.0990***
β_{time}	0.2019	-0.7017***
$\beta_{D_{2004-2005}}$	-0.0401	0.2442***
$\beta_{D_{2005-2006}}$	-0.2391**	0.5304***
γ	0.0012***	0.0026***
ρ	0.3396***	0.4860***
GARCH parameters		
ϖ	0.0151***	0.0315***
δ	0.9413***	0.8691***
φ	0.0254	0.0257
Standard deviations		
σ_{μ}	2.2147***	8.4446***
σ_{υ}	6.3419***	2.0899***
Probabilities (in %)		
λ_{I}	48.5613***	41.5957***
λ_{II}	41.1580***	50.4989***
λ_{III_a}	2.4462***	1.6899***
λ_{III_b}	0.0000	0.9188**
λ_{IV}	2.8750***	0.4921
λ_{V}	0.0028	3.0477***
λ_{VI}	4.9568***	1.7569
Log likelihood	-982.6623	-991.7400
LR tests		
H_0: $\gamma = 0$	128.868 *(0.000)*	115.345 *(0.000)*
H_0: $\sigma_{\mu} = \sigma_{\upsilon}$	44.278 *(0.000)*	41.018 *(0.000)*
H_0: $\rho = \delta = \varphi = 0$	206.992 *(0.000)*	228.932 *(0.000)*
Observations	723	723

Note: Asterisks indicate significance at 0.10*, 0.05** and 0.01*** levels, respectively. Numbers in parentheses are the *p*-values of the χ^2 statistics.

Fifth, the estimated autocorrelation coefficients ρ are significant at the 0.01 level. These estimated values are positive and their order of magnitude is comparable to those of the series R_{jit}.

Sixth, the estimated ARCH coefficients δ are highly significant in both directions, which indicates that the variance of the errors term ε_{jit} has a time-varying nature. Moreover, the estimated values for δ are high, which suggests that in both directions the variance of the marginal profit to spatial arbitrage obtained in regimes I & II is directly affected by the preceding shocks. Consistent with the large empirical literature dedicated to the dynamics of natural gas markets, these high values suggest the presence of a "volatility clustering" phenomenon.

Seventh, the null hypothesis of a static model (H_0: $\rho = \delta = \varphi = 0$) is firmly rejected by the data. This finding clearly justifies the need to use a dynamic specification to correct for the presence of both autocorrelation and time-varying heteroscedasticity.

Lastly, note that the use of a unique variance parameter for the half-normal distributions of the two non-negative error terms μ_{ijt} related to regimes III_a, III_b, and IV and υ_{ijt} related to regimes V and VI is firmly rejected by the data.

4.5 Model validation

The specification above is based upon arbitrary distributional assumptions of normal and half-normal errors and these distributional choices are, of course, questionable. According to the Monte Carlo experiments reported by Barrett and Li (2002), deviations from the half-normal distribution for the positive errors μ_{jit} and υ_{jit} are not problematic. In contrast, the use of normal assumption for both regimes I and II deserves some investigations to verify the validity of the PBM approach.

So far, little attention has been paid to these verifications in the PBM literature. To explore this question, we propose using the obtained estimates to generate, for each trade direction, two time series of binary indicator variables $\hat{d}^{I}_{jit}$ and $\hat{d}^{II}_{jit}$ indicating whether or not regime I, respectively II, is the regime with the highest probability at time t. The construction of these semiparametric estimates of time-varying regime probabilities follows those detailed in Barrett and Li (2002):

If $A_{jit} = 0$ then $\hat{d}^{I}_{jit} = 0$ and

$$\hat{d}^{II}_{jit} = 1 \quad \text{if} \quad \lambda_{II} f^{II}_{jit}\left(\eta_{jit} \middle| \eta_{ji(t-1)}, \theta_2\right) > \max\left(\lambda_{IV} f^{IV}_{jit}\left(\eta_{jit} \middle| \eta_{ji(t-1)}, \theta_2\right), \lambda_{VI} f^{VI}_{jit}\left(\eta_{jit} \middle| \eta_{ji(t-1)}, \theta_2\right)\right)$$

$$= 0 \quad \text{otherwise}$$

If $A_{jit} = 1$ then $\hat{d}^{II}_{jit} = 0$ and

$$\hat{d}^{I}_{jit} = 1 \quad \text{if} \quad \lambda_{I} f^{I}_{jit}\left(\eta_{jit} \middle| \eta_{ji(t-1)}, \theta_2\right) > \max\left(\left(\lambda_{III_a} + \lambda_{III_b}\right) f^{III}_{jit}\left(\eta_{jit} \middle| \eta_{ji(t-1)}, \theta_2\right), \lambda_{V} f^{V}_{jit}\left(\eta_{jit} \middle| \eta_{ji(t-1)}, \theta_2\right)\right)$$

$$= 0 \quad \text{otherwise}$$

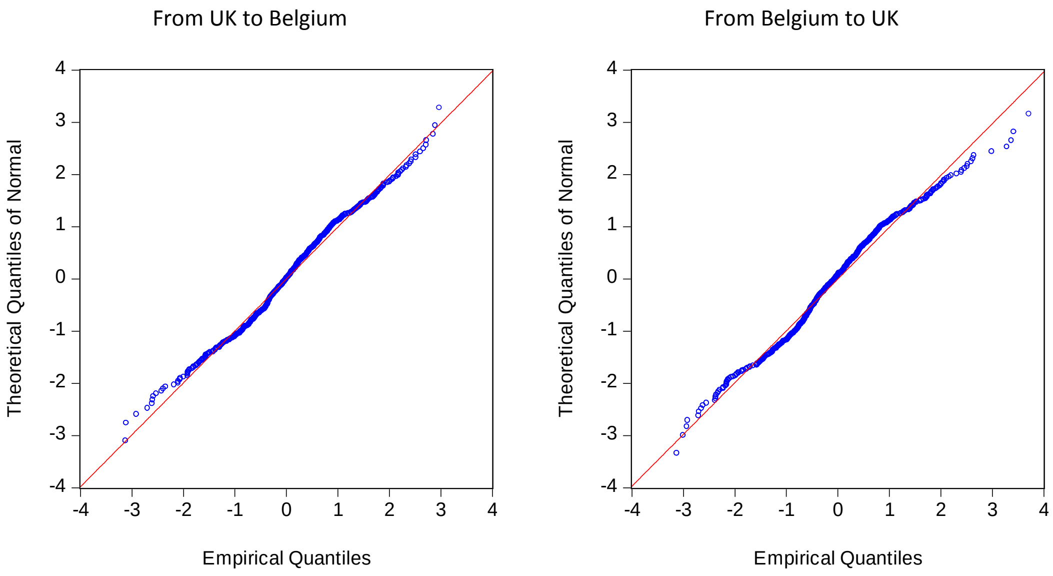
Assuming that the regime with the highest probability is the one that generated each observation, we can gather all the observations in the sample period that verify $\hat{d}^{I}_{jit} + \hat{d}^{II}_{jit} = 1$ and examine the distributional properties of the standardized residual series $\hat{e}_{jit} = \hat{\varepsilon}_{jit} / h_{jit}$.

Table 5. Descriptive statistics for the standardized residual series (sample: $\hat{d}^{I}_{jit} + \hat{d}^{II}_{jit} = 1$)

	From UK to Belgium	From Belgium to UK
Mean	0.086	-0.092
Median	0.079	-0.143
Maximum	2.973	3.713
Minimum	-3.122	-3.124
Std. Dev.	1.006	1.021
Skewness	-0.017	0.230
Kurtosis	3.556	3.981
Observations	650	677

Table 5 reports the descriptive statistics for the standardized residual series. These figures indicate the presence of both mild asymmetry and leptokurtosis. To gain further insights on the magnitude of these departures from normality, Figure 2 illustrates the quantile-quantile (Q-Q) plots of the standardized residual (i.e., the empirical quantiles versus the quantiles of a standard normal distribution). Each of these two Q-Q plots indicates a relatively good fit between the empirical quantiles and the theoretical one, except for several outliers. According to the Monte Carlo experiments reported in Barrett and Li (2002), moderate departures from normality caused by both mild asymmetry and leptokurtosis are not really problematic. Thus we proceed, assuming the relevance of the normal distributional assumption.

Figure 2. Q-Q plots of the standardized residual series (sample: $\hat{d}^{I}_{jit} + \hat{d}^{II}_{jit} = 1$)



5. Concluding remarks

The question of how to detect market power in the spatial arbitrages observed in a restructured natural gas industry is one of the key challenges that regulators and competition authorities across the world have to address. The objective of this paper is to offer an empirical methodology which is able to test for the presence of perfect competition in these spatial arbitrages. Our approach explicitly builds upon the literature dedicated to natural gas markets integration and extends it by focusing on the relationship between the observed spatial price difference and the intermarket trade flows.

A case study focusing on the IUK pipeline during the period 2003–2006 provided us with an opportunity to obtain a series of original findings. The estimated probability of spatial market equilibrium conditions holding is very high, suggesting high degrees of wholesale natural gas market integration, consistent with previous research on IUK price co-movements (Neumann et al., 2006). But, the empirical evidence also suggests the presence of imperfect competition in the observed spatial arbitrages, consistent with the price-data results in Rupérez-Micola and Bunn (2007). Although our discussion is centered on this specific infrastructure, it should be clear these results imply that some care is needed when interpreting the high degree of co-movements which is typically documented in the empirical studies conducted on European spatial market price data. Though these co-movements can be interpreted as objective signs of market integration, they do not necessarily reveal the existence of a perfectly competitive internal market.

Experience indicates that the institutional arrangements implemented in the UK have played a large part in shaping the EU's restructuration process. Future research will thus examine whether or not market equilibrium conditions hold in less mature continental markets. Such research could be useful for informing the current EU regulatory debates related to the functioning of the internal market for natural gas.

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