Title page

Does multimarket contact affect prices in the retail fuel industry? First empirical evidence

Jacint Balaguer*[a], Jordi Ripollés[a]

[a] *Department of Economics and Institute of International Economics. Universitat Jaume I Av. de Vicent Sos Baynat, s/n. 12071 Castelló, Spain.*

* Corresponding author

Email addresses: jacint.balaguer@uji.es (Jacint Balaguer*), jripolle@uji.es (Jordi Ripollés)

Preprint submitted to Energy Economics *October 22, 2020*

Abstract

Although multimarket competition is a fairly common characteristic in the retail fuel industry, its effect on the corresponding prices remains unexplored. By using a large dataset of service stations in Spain, a first set of empirical results for this industry is provided. It is consistently found that multimarket contact among chain stations increases prices, which adds support to the classical mutual forbearance hypothesis. However, it is further revealed that such an effect is critically dependent on the degree of concentration of each local market. Since the impact becomes more relevant in markets where concentration is lower, it also seems advisable to closely monitor these cases. Findings also suggest that, in those models where multimarket contact information is ignored, the specific effect of local concentration on prices could be significantly underestimated.

Keywords: Multimarket contact, spatial competition, prices, retail fuel industry.

JEL codes: D43, L41, L81, Q41.

1. Introduction

When firms meet others in several markets, it is possible that each of them resorts to a live-and-let-live strategy, which would cause higher prices than in a situation of rivalry in a single market. Since this hypothesis was formally proposed in Edwards (1955), a growing body of empirical work has been devoted to testing the occurrence of this phenomenon for a wide variety of industries, such as movie theaters (Feinberg, 2014), leasing (Degl'Innocenti et al., 2014), newspapers (Fu, 2003), cement (Jans and Rosenbaum, 1997), retail grocery (Aalto-Setälä, 2002), hospitals (Schmitt, 2018), mobile telephones (Parker and Röller, 1997; Busse, 2000), hotel accommodation (Fernandez and Marin, 1998; Silva, 2015), retail banking (Pilloff, 1999; De Bonis and Ferrando, 2000; Coccorese and Pellecchia, 2013; Kasman and Kasman, 2016) or airlines (Evans and Kessides, 1994; Gimeno and Woo, 1999; Bilotkach, 2011; Zou et al., 2012; Ciliberto and Williams, 2014; Ma et al., 2020). To our knowledge, there is no empirical evidence for the retail fuel industry, despite the fact that the typical large networks in this sector mean that companies generally meet in different geographical markets. This paper addresses this research gap by using a panel dataset from the population of petrol stations on the Spanish peninsula.

Results from this paper can be especially interesting for researchers and policy makers concerned with competition in this industry. On the one hand, the results about the impact of multimarket contact on prices can help to design better sectoral policies. In fact, regardless the concentration restriction measures, monitoring multimarket contact could be an additional tool improving effective competition in the sector. This could be particularly relevant at least for Spain, where the regulatory agency has repeatedly expressed concern about prices being relatively higher than in neighboring countries (e.g., CNMC, 2018). On the other hand, including an indicator of multimarket contact in the competition analysis can be seen as a natural extension of those models devoted to measuring the impact of the local concentration of stations in home markets on fuel prices (e.g., Nowakowski and Karasiewicz, 2016; Tappata and Yan, 2017; Bernardo, 2018; Oczkowski et al., 2018; Kvasnička et al., 2018; Balaguer and Ripollés, 2020).[1] Considering contacts in non-home markets could imply more accurate estimates of the concentration effect if contacts are useful to explain prices

[1] An interesting survey on a first generation of papers on this topic can be seen in Eckert (2013).

and, in turn, they are correlated with local concentration. While the first condition is the subject of this study, the existence of correlation seems to be quite plausible. That is, insofar as many companies operate in a same territory, we would expect a lower number of contacts in the case that markets were more concentrated.

The rest of this paper is organized as follows. In the next section, we will offer a brief overview of the background to this work. In Section 3, we will describe our data and the procedure used to build the required variables. In Section 4, we will specify the empirical model, make some methodological considerations for their estimation, and discuss the results obtained from regressions. Tests for robustness using alternative measures of industrial concentration are carried out in Section 5. Finally, in Section 6, we will draw some conclusions and policy implications.

2. Background

How spatial competition in the home markets affects prices has been one of the central questions in the field of industrial organization for some time. The practical application of the answer to this inquiry is probably one of the main reasons for the existence of a significant number of empirical studies on industries monitored by antitrust and regulation agencies. As a result, we currently benefit from a relevant amount of work focused on exploring the relationship between spatial competition and price levels in the retail fuel industry.

As can be inferred from Eckert (2013), as well as from a more current review of this literature, improving the measurement of the specific effect of spatial competition has been one of the essential concerns in this area of research. It has been broadly acknowledged that the ability to control other factors affecting prices contributes to this purpose. In fact, a great effort has been made to collect station-specific characteristics. For example, variables on brand (e.g., Barron et al., 2004; Firgo et al., 2015), distance to refinery (e.g., Pennerstorfer, 2009; Haucap et al., 2017), population density (e.g., Marvel, 1978; Hosken et al., 2008) or traffic intensity index (e.g., Firgo et al., 2015; Haucap et al., 2017) have been introduced into the empirical models as control variables. The control of a wide set of the unobserved characteristics has been facilitated in another generation of works using panel data models (e.g., Hastings, 2004; Lach and

Moraga-González, 2017; Balaguer and Ripollés, 2020), which in turn have also led to an improvement in the efficiency of the estimators due to the usual increase in degrees of freedom and sample variability.

The way the relevant market is defined is another of the critical aspects where substantial advances have also been carried out. The empirical literature offers us, on the one hand, a generation of studies where the local competition, calculated by a simple count of the number of competitors or by concentration, has been defined for different gross administrative areas (e.g., Marvel, 1978; Van Meerbeeck, 2003; Sen, 2005; Clemenz and Gugler, 2009). For example, the early paper by Marvel (1978) estimated the impact on retail gasoline prices from the concentration index in the years 1964-1971, taking as relevant markets each of the 22 major U.S. cities. The convenience of defining markets from the point of view of the specific degree of competition that each station faces has been gaining ground among researchers. So, we already have a significant number of analyses incorporating this idea. Inspired by a paper by Shepard (1991), it was initially materialized by drawing circles centered on each station, where a radius of a reasonable predefined length sought to guarantee that the sellers within it can be perceived as substitutes for consumers (e.g., Barron et al., 2004, Hosken et al., 2008; Pennerstorfer, 2009; Albalate and Perdiguero, 2015). Indeed, the influential paper by Barron et al. (2004) tried to capture the degree of spatial competition faced by each station by counting the number of rivals existing within a circle with a radius of 1.5 miles centered on each of them. In order to capture the relevant rivals in a more realistic way than by computing a simple Euclidean distance, a relevant part of recent work has alternatively used a more accurate strategy based on the travel routes offered by geo-information technologies. These methods take into account the rivals within a certain travel-distance through the different routes that depart from each station (e.g., Tappata and Yan, 2017; Kvasnička et al., 2018). Similarly, other studies have taken into consideration the rivals that can be reached during a certain travel-time (e.g., Perdiguero and Borrell, 2019; Balaguer and Ripollés, 2020). A notable paper in this respect is the one by Perdiguero and Borrell (2019) for the stations operating in Catalonia (Spain). By adopting this last approach, it provides an especially interesting conclusion for our purpose. The author's findings indicate that, to better capture the relevant rivals, local markets should be defined through a 5- to 6-min travel-time isochrone around each station.

Regardless of the degree of precision in defining markets, literature shows us that prices are generally higher when the density of competitors is lower (e.g., Barron et al., 2004; Cooper and Jones, 2007; Tappata and Yan, 2017) or, in papers considering concentration, when this last alternative metric is greater (e.g., Marvel, 1978; Sen and Townley, 2010; Nowakowski and Karasiewicz, 2016). Because results depend on the context being analyzed, let us focus on the particular case of a paper by Bernardo (2018) carried out for the Metropolitan Area of Barcelona (Spain). In that study it is suggested that the entry of a new competitor within a one-mile radius of a station hardly implies any reduction in its prices. This outcome, which calls into question the success of a set of measures implemented in 2013 with the aim of promoting entries of competitors in the Spanish territory, is complemented by a recent paper by Balaguer and Ripollés (2020). This last study shows us that a more significant price drop would be achieved through the entry of stations associated with supermarkets or advertised as "low cost".

One of the issues that, to our knowledge, has not yet been studied is whether retail fuel prices could be further affected by the contacts in non-home markets. Given that this industry is largely characterized by firms operating in multiple local markets, it would be plausible to think that filling this gap might be important to avoid problems of omission of relevant variables in the models as well as to help to design better competition measures in this sector. At this point, it seems reasonable to ask ourselves how these contacts might affect retail fuel prices. The classic answer is obtained from the seminal paper by Edwards (1955), which indicates that firms that meet in several markets may be tempted to collude because the prospects of gains in a specific local market do not make it worthwhile running the risk of a generalized rivalry. In order to empirically validate this hypothesis of mutual forbearance, we can see considerable concern in the literature about employing appropriate multimarket contact indicators. In fact, it is not surprising to find studies that use more than a single indicator to alternatively capture the effect of these contacts (e.g., Jans and Rosenbaum, 1997; Coccorese and Pellecchia, 2013; Pham et al., 2020). Results are usually robust to the indicators used, offering quite a lot of support for the mutual forbearance hypothesis in a variety of industries, including such highly studied sectors like retail banking (e.g., Pilloff, 1999; Coccorese and Pellecchia, 2009; Coccorese and Pellecchia, 2013; Molnar et al., 2013; Pham et al., 2020) or airlines (e.g., Evans and Kessides, 1994; Bilotkach, 2011; Zou et al., 2012; Ciliberto and Williams, 2014; Ciliberto et al., 2019; Ma et al., 2020).

Although it seems obvious that the impact of multimarket contact on prices depends on the industry analyzed, it might appear less evident that it could be critically conditioned by each local market structure in the same industry. However, as Bernheim and Whinston (1990) demonstrated, we currently know that the marginal response of prices to contacts in non-home markets can be negatively dependent on the degree of concentration in the home market, so that the support for the classical mutual forbearance hypothesis can be weakened or even refuted in those highly concentrated markets. According to the authors we have just mentioned, this is because companies subject to these contacts could transfer power from highly concentrated markets to the more competitive ones. It should be noted that, even before this last theoretical work we have just cited, this interesting outcome received empirical support in the literature (Mester, 1987; Fernandez and Marin, 1998, Hannan, 2006; Coccorese and Pellecchia, 2013).

3. Data and variables

The dataset employed in this article comes from the Hydrocarbons Geoportal (www.geoportalgasolineras.es) of the Spanish Ministry for the Ecological Transition. This website provides real-time information about retail diesel prices, geographical coordinates and brands concerning every gas station located in the Spanish peninsula. For the sake of confidentiality, historical data series are not made available. So, we have assembled a quarterly dataset by periodically collecting information from the above mentioned website over the period from January 1, 2011, to June 30, 2016. The result is an unbalanced panel dataset of 22 quarters, the number of individual observations of which range from 7,688 to 10,876 due to permanent or temporary closures of stations throughout the period. Although some of these stations are independent from a network, an important number of them (ranging from 6,098 to 6,945) belong to a certain firm with at least two stations (hereafter, "chain").[2]

[2]The number of chains ranges from 69 to 102, depending on the quarter considered.

To measure the importance of spatial competition and multimarket contact we need to define the relevant markets. In accordance with recent research on spatial competition, we decided to delineate certain driving-time isochrones surrounding each station. We chose a 5-minute isochrone in line with the recommendation for this industry in Spain derived from Perdiguero and Borrell (2019). The practical implementation was performed by considering the optimal vehicle route for each pair of sampled stations, according to the Open Source Routing Machine service which, in turn, is based on the road network and speed limits (sourced from OpenStreetMap).

The degree of local market concentration facing each station i in each quarter t was approximated by a pseudo Herfindahl-Hirschman index. It is noted as $C_{it}^{HHI} = \left(\sum_{k=1}^{K} (n_k/N)^2\right)_{it}$, where n_k is the number of local stations associated to firm k and N represents the total number of local stations. In order to apply this index, we first identify those firms consisting of an independent station and those composed of stations belonging to a chain. Hence, for each delineated local market, we calculated the relative presence of each firm. That is, $(n_k/N) = (1/N)$ for the independent stations, and $(n_k/N) \geq (1/N)$ in the case of chains. The index C_{it}^{HHI} ranges from $(1/N)$ to 1, taking low values when local markets are characterized by a large number of stations associated to firms with homogeneous relative presences, and achieves the maximum level of 1 in the case of a local monopoly.

Two measures for multimarket contact were computed following the standard empirical literature (e.g., Evans and Kessides, 1994; Jans and Rosenbaum, 1997; Coccorese and Pellecchia, 2013). We first calculated a simple indicator MM_{it}^{I} as the average number of non-home market contacts (expressed in thousands) for each pair of firms competing within the home market where station i operates. Even though this indicator is easily comprehensible, it does not acknowledge that firms' incentives to collude might depend on the relative importance of each contact counterpart. Therefore, we have also considered a more sophisticated measure, denoted by MM_{it}^{II}, which further takes into account this last fact. Specifically, it has been computed as the average sum of relative presences in non-home markets per contact generated in these markets, regarding the firms competing within the home market where station i operates. The details on how MM_{it}^{I} and MM_{it}^{II} are calculated can be seen in Appendix A. Both measures are set to zero for stations associated to a firm without home contacts (i.e., local monopolies) and without non-home contacts (i.e., stations independent from a network or chains operating exclusively in the home market). In contrast, the measures take higher values insofar as the stations belong to chains that meet with independent brands or other chains in more markets, attaining their maximum value when such meetings occur in all possible local markets.

Table 1 offers an overview of the average level of the main variables (with their standard deviation) for the entire period, as well as for sub-periods, which gives us a general idea of their evolution. Retail prices (P_{it}) have been expressed net of taxes and, as is well known, their evolution is fundamentally determined by changes in wholesale oil prices in international markets. Indeed, the retail price decrease from 2014 as a consequence of a noticeable and persistent drop in international oil prices. Although the retail prices have been decreasing in the last part of the sample, they are still relatively higher than most of the European Union countries (e.g., CNMC, 2018). Examining if business concentration and/or multimarket contacts are contributing significantly to this peculiarity is part of this work. From the summary statistics, we can also see that the concentration in local markets has progressively decreased, which is compatible with the plausible relocation of some stations in places where others operate, but also with a regular increase in the total number of stations in Spain. This last fact is easily verifiable from our database and consistent with the aggregate information provided by the annual reports of the Spanish Association of Operators of Petroleum Products (AOP).[3] So, for example, these reports reveal that the inter-annual average increase in the number of stations was about 0.9% from December 2010 to December 2012, while it amounted to an inter-annual average growth of 1.78% from then until December 2016. This stronger growth in the latter sub-period was a consequence of the appearance of new operators belonging to supermarkets and basically derived from an extraordinary entry of independent sellers. A regulation for the industry implemented in early 2013 (Royal Decree-Law 4/2013) was the one that fostered the entry of these types of stations lacking large networks. This may be the main reason why, in general, we can observe in Table 1 that there is a drop in multimarket measures from 2012.

[3]These reports are collected on the AOP website (https://www.aop.es).

[Please insert Table 1 about here]

Correlations among the variables are shown at the end of Table 1. Retail prices are positively correlated with both local market concentration and the measures of multimarket contact. In addition, these contact measures, MM_{it}^{I} and MM_{it}^{II}, are negatively correlated with local market concentration C_{it}^{HHI}. As expected, higher (lower) rates of market concentration are associated with less (more) multimarket contacts. Therefore, under this context, if contacts in non-home markets turn out to be a relevant factor in determining home prices, their omission in the regression analysis is likely to lead to a significant understatement of the effect of local market concentration on prices. This will be studied later on.

4. Specification and estimation

As has been outlined in Section 2, a broad body of literature on retail fuel focuses on local market concentration (or density) as one of the main variables responsible for prices in this industry. In order to know whether multimarket contact has some relevance in the retail pricing behavior, in the present analysis we extend the approach underlying this type of works. For this purpose, we specify the following regression model:

$$lnP_{it} = \beta_1 C_{it} + \beta_2 MM_{it} + \beta_3 C_{it} MM_{it} + \theta_t + \lambda_z + \lambda_b + u_{it} \tag{1}$$

where the natural logarithm of prices (net of taxes) set by each station i at each time t (lnP_{it}) is expressed as a linear function of the degree of concentration in the home local market (C_{it}), and a measure of multimarket contact (MM_{it}). An interaction term ($C_{it}MM_{it}$) is also considered in line with the models empirically initiated by Mester (1987) and later taken into account in theoretical approaches (Bernheim and Whinston, 1990). So, we further acknowledge that differences in concentration levels may imply variations in the price effect of non-home market contacts. We have also introduced time fixed effects (θ_t) to control for common cost variations (e.g., regular wholesale price changes), as well as postcode (λ_z) and brand effects (λ_b) to control for neighborhood (e.g., income) and firm idiosyncrasies (e.g., consumer perception of quality), respectively. Lastly, u_{it} is the error term.

Estimating Equation (1) by Ordinary Least Squares (OLS) could lead to endogeneity problems due to the potential presence of simultaneity between prices, market concentration, and multimarket contacts. Therefore, besides using OLS, we also apply the Two Stages Least Squares (2SLS) procedure in order to address this concern. As usual (e.g., Evans et al., 1993; Evans and Kessides, 1994; Reed, 2015), by exploiting the panel data structure, we instrument the measures of multimarket contact and market concentration by their own lagged values. Additionally, as in other papers on spatial competition (e.g., Sen and Townley, 2010; Clemenz and Gugler, 2009), the municipality-specific population (in our case collected from the Spanish Statistics Office) has also been included as an instrumental variable.

Table 2 provides the OLS and 2SLS estimates, using the Driscoll and Kraay (1998) standard errors robust to heteroskedasticity, serial correlation, and cross-sectional dependence.[4] We also report some diagnostic tests which, at the standard statistical significance levels, support the adequacy of the instruments employed and the 2SLS estimator, both implementing MM_{it}^{I} and MM_{it}^{II}. More specifically, the Hansen J test indicates that the instruments are exogenous, the Kleibergen-Paap F test supports a strong correlation between instruments and explanatory variables, and the Kleibergen-Paap LM test indicates that the model is not under-identified. Lastly, the Durbin-Wu-Hausman test confirms the presence of endogeneity, endorsing the use of 2SLS rather than OLS.

[Please insert Table 2 about here]

In view of the diagnostic tests just commented on, we focus our attention on the results from the 2SLS procedure. In column I we show the outcomes by estimating a constrained version of Equation (1), where multimarket contact information is excluded. The positive sign for the concentration coefficient is widely consistent with the research for the retail fuel sector both in Spain (i.e., Bernardo, 2018; Balaguer and Ripollés, 2020) and in other countries (e.g., Barron et al., 2004; Clemenz and Gugler,

[4]These problems have been revealed by using the Greene (2003) test for groupwise heteroskedasticity, the Wooldridge (2010) test for serial correlation, and the Pesaran (2004) test for cross-sectional dependence. In all cases, the corresponding null hypotheses can be rejected with p-values lower than 1%.

2009; Sen and Townley, 2010; Nowakowski and Karasiewicz, 2016; Tappata and Yan, 2017; Oczkowski et al., 2018). The rationality of this effect lies in that the more concentrated retailers' local environment is, the higher they are able to set their prices.

In order to test the mutual forbearance hypothesis of Edwards (1955) in the industry, let us now focus our attention on the regressions where the multimarket contact information is considered. As can be seen, our results indicate that retail pricing behavior is significantly dependent from both the importance of non-home contacts and their interaction with the market concentration. The signs of the estimated coefficients are clearly consistent, regardless of whether MM_{it}^{I} or MM_{it}^{II} is used. However, in order to know as accurately as possible the marginal effect of multimarket contact on prices, we select the most reliable estimations according to the Akaike information criterion (AIC). Results from the AIC indicate that the further information introduced to build the MM_{it}^{II} measure seems particularly suitable to improve the quality of the model. How (the log of) prices are marginally affected by multimarket contact according to this last measure is shown in Figure 1. As we can see, since the marginal responses of prices to multimarket contact are statistically positive regardless of market concentration, the mutual forbearance hypothesis is clearly supported. That is, those stations for which the contacts are more relevant set higher prices, and vice versa. This finding seems quite consistent with a more aggressive pricing strategy in Spain from a group of stations that are independent or associated to very limited networks. We are specifically referring to those retailers associated with the supermarkets and those labeled as "low cost" (i.e., Bernardo et al., 2014; Balaguer and Ripollés, 2020).[5]

[Please insert Figure 1 about here]

Let us now focus on the specific effect of the interaction term ($C_{it}MM_{it}$). In line with previous papers based on data for other industries, such as hotels and banks (e.g., Fernandez and Marin, 1998; Hannan, 2006; Coccorese and Pellecchia, 2013), the marginal response of prices to multimarket contact is negatively dependent on the degree of local market concentration. That is, as we can see in Figure 1, in very

[5]According to our sample, while stations of supermarkets and those known as "low cost" exhibit an average of 42 non-home contacts, the remaining stations present 300 contacts.

concentrated markets the external contacts play a less relevant role. The degree of concentration is, in this last case, enough to establish higher prices. The opposite would happen under situations of lower concentration. The mutual forbearance offsets, to some extent, the drop in prices in those home markets with highly competitive structures.

Before concluding this section, we ask ourselves if the estimated coefficient of concentration tends to be biased when multimarket contact information is omitted. For this purpose, we can compare the estimated effect of concentration on prices derived from the constrained model with respect to that obtained in the selected extended model, where such an effect is further dependent on the importance of the non-home contacts. From a comparison, at the sample average of MM_{it}^{II} that can be seen in Table 1 (i.e., 0.315), we obtain that the impact is clearly underestimated. That is to say, the marginal effect in the constrained model (0.816×10^{-2}) is about 50% lower than in the preferred model (1.65×10^{-2}) at the average level of MM_{it}^{II}.

5. Robustness check

To test the sensibility of our results to alternative measures of local market concentration, we now consider three indicators inspired by the proposals of Hall and Tideman (1967), Rosenbluth (1961), and Horvath (1970). To apply these indicators, firms in each market are ordered according to their relative presence (n_k/N), where $k = 1$ is the firm with the highest presence and $k = K$ is the one with the lowest. The Hall-Tideman index is then calculated as $C_{it}^{HTI} = (2\sum_{k=1}^{K}(k \cdot (n_k/N)) - 1)^{-1}$. The Rosenbluth index is constructed in an analogous manner to the previous one, but (n_k/N) is weighted by the reverse order of k. That is, $C_{it}^{RI} = (2\sum_{k=1}^{K}((K - k + 1) \cdot (n_k/N)) - 1)^{-1}$. Finally, the Horvath index, also known as the comprehensive concentration index, is defined as $C_{it}^{CCI} = (n_1/N) + \sum_{k=2}^{K}(n_k/N)^2(2 - (n_k/N))$. It should be noted that these alternative indexes range from $(1/N)$ to 1, where low values indicate competition and the maximum value of 1 captures a local monopoly.

[Please insert Table 3 about here]

The constrained and unconstrained versions of Equation (1) have been estimated again, now considering the indicators of market concentration described above. For the sake of simplicity, in Table 3 we display the results from using the multimarket contact measure MM_{it}^{II}, which is preferable according to the AIC. As can be seen, in all cases the diagnostic tests confirm the validity of the instruments employed, the identification of the model, and the adequacy of the 2SLS over the OLS estimator. Although the coefficients are not directly comparable because they use different concentration indexes, we note that they are consistent with those previously obtained. Specifically, we find that the marginal effect of multimarket contact on retail prices is significantly positive, but this depends negatively on the degree of local market concentration.

Finally, we are also interested in knowing whether the coefficient of concentration tends to be biased in the constrained version of Equation (1). As can be seen in Table 3, underestimation is once again revealed, and this outcome is clearly robust to the different measures of concentration used here. Indeed, at the sample average of MM_{it}^{II}, the estimated marginal effect of concentration is about 63%, 74% and 40% lower than when multimarket contact is introduced in the model by using C_{it}^{HTI}, C_{it}^{RI} and C_{it}^{CCI} respectively. This result, together with the one in the preceding section, raises a serious concern about the possibility of the relationship between the market structure and prices being undervalued in decision-making based on previous estimates from literature. We hope that future research provides further evidence in order to add greater clarity to this important question.

6. Conclusion and policy implications

The analysis carried out here constituted a straightforward extension of the standard models concerned with competition in retail fuel. It provides the industrial organization literature with the first attempt to examine the mutual forbearance hypothesis in the industry. We have used quarterly data for a period of more than five years on the set of stations operating in the Spanish peninsula, which offers us a relatively high number of observations compared with other analyses on this topic. Our results may be especially interesting to help future policy design, as well as to improve the research on competition in the retail fuel sector.

We consistently found that multimarket contact between chain stations increases prices, which adds support to the hypothesis indicating that mutual forbearance derived from these meetings relaxes competition (Edwards, 1955). This result suggests that, in contexts where contacts are relevant, it will be more difficult to achieve a high degree of effective competition. In this sense, it seems reasonable to promote the relative presence of those stations independent from chains. Moreover, regulators must be careful with the authorization of mergers for companies with large chains even though they operate across different parts of the territory. Although the merger did not involve a general increase in the concentration in local markets, it can lead to higher prices than under the original situation. We also found that the effect of those contacts is dependent on the degree of local market concentration. More specifically, it becomes weaker (stronger) in those markets where concentration is higher (lower). This finding highlights an interesting business strategy, namely, contacts between chain stations lose some relevance where it is easier to achieve collusive outcomes. Conversely, in more competitive markets, the multimarket linkages would be providing chains with an effective coordination mechanism to achieve certain power and, consequently, increase prices. Therefore, even under apparently competitive markets, it seems advisable to carefully monitor them if stations associated with large chains operate within them.

Another key point we learned is that it seems advisable to further introduce multimarket contact information in the typical models devoted to knowing how the effective competition of stations is affected by concentration. In fact, it is revealed that the coefficient associated to concentration will tend to be underestimated when indicators on these contacts are ignored. Policies aimed at restricting the presence of large companies or facilitating the entry of independent stations could be, then, undervalued. This could explain, to some extent, why in part of the literature it is suggested that reducing concentration has a virtually insignificant effect on retail fuel prices.

We hope that this first analysis promotes future research on deeper aspects related to the multimarket contact effect in retail fuel prices. Given the particular characteristics of the industry, such as the relative variability of its input costs or the typical vertical integration in some of the companies, it could be interesting to know the importance of the the mutual forbearance hypothesis under different cost levels or types of companies. Notwithstanding deeper studies can be conducted, we believe that competition policies in different countries could benefit from considering the proposed model extension. Until then, when competition policies are applied in the sector, it seems necessary to take into account that the models based on the typical approach could be seriously undervaluing the effect of local market structure on retail fuel prices.

Acknowledgements

The authors thank the anonymous referees for their valuable remarks and suggestions on the earlier version of the paper. The authors alone are responsible for this work and any shortcomings that may remain. Financial support from the Spanish Ministry of Science, Innovation and Universities (RTI2018-100983-B-I00) and from Universitat Jaume I (UJI-B2019-53) is also gratefully acknowledged.

References

Aalto-Setälä, V. (2002). The effect of concentration and market power on food prices: evidence from finland. *Journal of Retailing*, 78(3):207–216.

Albalate, D. and Perdiguero, J. (2015). Entry regulation asymmetries and petrol competition in a mixed motorway network. *Journal of Transport Economics and Policy (JTEP)*, 49(4):603–625.

Balaguer, J. and Ripollés, J. (2020). Do classes of gas stations contribute differently to fuel prices? Evidence to foster effective competition in spain. *Energy Policy*, 139:111315.

Barron, J. M., Taylor, B. A., and Umbeck, J. R. (2004). Number of sellers, average prices, and price dispersion. *International Journal of Industrial Organization*, 22(8-9):1041–1066.

Bernardo, V. (2018). The effect of entry restrictions on price: evidence from the retail gasoline market. *Journal of Regulatory Economics*, 53(1):75–99.

Bernardo, V., González, J. L. J., and Perdiguero, J. (2014). Competencia y medidas liberalizadoras en el sector minorista de hidrocarburos: análisis del efecto sobre los precios de la entrada de nuevos operadores en españa. *Cuadernos Económicos de ICE*, (88):127–144.

Bernheim, B. D. and Whinston, M. D. (1990). Multimarket contact and collusive behavior. *The RAND Journal of Economics*, pages 1–26.

Bilotkach, V. (2011). Multimarket contact and intensity of competition: evidence from an airline merger. *Review of Industrial Organization*, 38(1):95–115.

Busse, M. R. (2000). Multimarket contact and price coordination in the cellular telephone industry. *Journal of Economics & Management Strategy*, 9(3):287–320.

Ciliberto, F., Watkins, E., and Williams, J. W. (2019). Collusive pricing patterns in the us airline industry. *International Journal of Industrial Organization*, 62:136–157.

Ciliberto, F. and Williams, J. W. (2014). Does multimarket contact facilitate tacit collusion? Inference on conduct parameters in the airline industry. *The RAND Journal of Economics*, 45(4):764–791.

Clemenz, G. and Gugler, K. (2009). Locational choice and price competition: some empirical results for the austrian retail gasoline market. In *Spatial Econometrics*, pages 223–244. Springer.

CNMC (2018). Supervision de la distribucion de carburantes en estaciones de servicio, Diciembre 2018. pages 1–28. Comision Nacional de Mercados y Competencia.

Coccorese, P. and Pellecchia, A. (2009). Multimarket contact and profitability in banking: evidence from italy. *Journal of Financial Services Research*, 35(3):245–271.

Coccorese, P. and Pellecchia, A. (2013). Multimarket contact, competition and pricing in banking. *Journal of International Money and Finance*, 37:187–214.

Cooper, T. E. and Jones, J. T. (2007). Asymmetric competition on commuter routes: the case of gasoline pricing. *Southern Economic Journal*, pages 483–504.

De Bonis, R. and Ferrando, A. (2000). The italian banking structure in the 1990s: testing the multimarket contact hypothesis. *Economic Notes*, 29(2):215–241.

Degl'Innocenti, M., Girardone, C., and Torluccio, G. (2014). Diversification, multimarket contacts and profits in the leasing industry. *Journal of International Financial Markets, Institutions and Money*, 31:231–252.

Driscoll, J. C. and Kraay, A. C. (1998). Consistent covariance matrix estimation with spatially dependent panel data. *Review of Economics and Statistics*, 80(4):549–560.

Eckert, A. (2013). Empirical studies of gasoline retailing: A guide to the literature. *Journal of Economic Surveys*, 27(1):140–166.

Edwards, C. D. (1955). Conglomerate bigness as a source of power. In *Business Concentration and Price Policy*, pages 331–359. Princeton University Press.

Evans, W. N., Froeb, L. M., and Werden, G. J. (1993). Endogeneity in the concentration–price relationship: Causes, consequences, and cures. *The Journal of Industrial Economics*, pages 431–438.

Evans, W. N. and Kessides, I. N. (1994). Living by the "golden rule": Multimarket contact in the us airline industry. *The Quarterly Journal of Economics*, 109(2):341–366.

Feinberg, R. M. (2014). Price effects of multimarket contact among movie chains in small us metropolitan areas. *Economics Letters*, 123(1):6–8.

Fernandez, N. and Marin, P. L. (1998). Market power and multimarket contact: Some evidence from the spanish hotel industry. *The Journal of Industrial Economics*, 46(3):301–315.

Firgo, M., Pennerstorfer, D., and Weiss, C. R. (2015). Centrality and pricing in spatially differentiated markets: The case of gasoline. *International Journal of Industrial Organization*, 40:81–90.

Fu, W. W. (2003). Multimarket contact of us newspaper chains: circulation competition and market coordination. *Information Economics and Policy*, 15(4):501–519.

Gimeno, J. and Woo, C. Y. (1999). Multimarket contact, economies of scope, and firm performance. *Academy of Management Journal*, 42(3):239–259.

Greene, W. H. (2003). *Econometric analysis*. Pearson Education India.

Hall, M. and Tideman, N. (1967). Measures of concentration. *Journal of the American Statistical Association*, 62(317):162–168.

Hannan, T. H. (2006). Retail deposit fees and multimarket banking. *Journal of Banking & Finance*, 30(9):2561–2578.

Hastings, J. S. (2004). Vertical relationships and competition in retail gasoline markets: Empirical evidence from contract changes in southern california. *American Economic Review*, 94(1):317–328.

Haucap, J., Heimeshoff, U., and Siekmann, M. (2017). Fuel prices and station heterogeneity on retail gasoline markets. *The Energy Journal*, 38(6).

Horvath, J. (1970). Suggestion for a comprehensive measure of concentration. *Southern Economic Journal*, pages 446–452.

Hosken, D. S., McMillan, R. S., and Taylor, C. T. (2008). Retail gasoline pricing: What do we know? *International Journal of Industrial Organization*, 26(6):1425–1436.

Jans, I. and Rosenbaum, D. I. (1997). Multimarket contact and pricing: Evidence from the us cement industry. *International Journal of Industrial Organization*, 15(3):391–412.

Kasman, S. and Kasman, A. (2016). Multimarket contact, market power and financial stability in the turkish banking industry. *Empirical Economics*, 50(2):361–382.

Kvasnička, M., Staněk, R., and Krčál, O. (2018). Is the retail gasoline market local or national? *Journal of Industry, Competition and Trade*, 18(1):47–58.

Lach, S. and Moraga-González, J. L. (2017). Asymmetric price effects of competition. *The Journal of Industrial Economics*, 65(4):767–803.

Ma, W., Wang, Q., Yang, H., and Zhang, Y. (2020). Is multimarket contact an antitrust concern? a case of china's airline market. *Transportation Research Part A: Policy and Practice*, 132:515–526.

Marvel, H. P. (1978). Competition and price levels in the retail gasoline market. *The Review of Economics and Statistics*, pages 252–258.

Mester, L. J. (1987). Multiple market contact between savings and loans: Note. *Journal of Money, Credit and Banking*, 19(4):538–549.

Molnar, J., Violi, R., and Zhou, X. (2013). Multimarket contact in italian retail banking: Competition and welfare. *International Journal of Industrial Organization*, 31(5):368–381.

Nowakowski, M. and Karasiewicz, G. (2016). Market structure and price-cost margins in european retail gasoline industry. *Journal of Management and Business Administration*, 24(3):105–124.

Oczkowski, E., Wong, A., and Sharma, K. (2018). The impact of major fuel retailers on regional new south wales petrol prices. *Economic Analysis and Policy*, 57:44–59.

Parker, P. M. and Röller, L.-H. (1997). Collusive conduct in duopolies: multimarket contact and cross-ownership in the mobile telephone industry. *The RAND Journal of Economics*, pages 304–322.

Pennerstorfer, D. (2009). Spatial price competition in retail gasoline markets: evidence from austria. *The Annals of Regional Science*, 43(1):133–158.

Perdiguero, J. and Borrell, J. R. (2019). Driving competition in local markets with near-perfect substitutes: an application on the spanish retail gasoline market. *Empirical Economics*, 57(1):345–364.

Pesaran, M. H. (2004). General diagnostic tests for cross section dependence in panels. *CESifo Working Paper Series, 1229*.

Pham, T., Talavera, O., and Yang, J. (2020). Multimarket competition and profitability: evidence from ukrainian banks. *Oxford Economic Papers*, 72(2):517–545.

Pilloff, S. J. (1999). Multimarket contact in banking. *Review of Industrial Organization*, 14(2):163–182.

Reed, W. R. (2015). On the practice of lagging variables to avoid simultaneity. *Oxford Bulletin of Economics and Statistics*, 77(6):897–905.

Rosenbluth, G. (1961). Address to" round table-gesprach uber messung der industriellen konzentration. In *Die Konzentration in der Wirtschaft*, volume 22, pages 391–394. Schriften des Vereins fur Socialpolitik NS.

Schmitt, M. (2018). Multimarket contact in the hospital industry. *American Economic Journal: Economic Policy*, 10(3):361–87.

Sen, A. (2005). Does increasing the market share of smaller firms result in lower prices? empirical evidence from the canadian retail gasoline industry. *Review of Industrial Organization*, 26(4):371–389.

Sen, A. and Townley, P. G. (2010). Estimating the impacts of outlet rationalization on retail prices, industry concentration, and sales: empirical evidence from canadian gasoline markets. *Journal of Economics & Management Strategy*, 19(3):605–633.

Shepard, A. (1991). Price discrimination and retail configuration. *Journal of Political Economy*, 99(1):30–53.

Silva, R. (2015). Multimarket contact, differentiation, and prices of chain hotels. *Tourism Management*, 48:305–315.

Tappata, M. and Yan, J. (2017). Competition in retail petrol markets. *Australian Economic Papers*, 56(3):201–218.

Van Meerbeeck, W. (2003). Competition and local market conditions on the belgian retail gasoline market. *De Economist*, 151(4):369–388.

Wooldridge, J. M. (2010). *Econometric analysis of cross section and panel data*. MIT press.

Zou, L., Yu, C., and Dresner, M. (2012). Multimarket contact, alliance membership, and prices in international airline markets. *Transportation Research Part E: Logistics and Transportation Review*, 48(2):555–565.

Table 1: Descriptive statistics

	P_{it}	C_{it}^{HHI}	MM_{it}^{I}	MM_{it}^{II}
Mean (standard deviation):				
Full period	0.990 (0.116)	0.613 (0.315)	0.246 (0.462)	0.315 (0.319)
2011	1.038 (0.023)	0.632 (0.311)	0.246 (0.464)	0.319 (0.330)
2012	1.096 (0.024)	0.626 (0.311)	0.258 (0.481)	0.319 (0.326)
2013	1.068 (0.029)	0.618 (0.313)	0.250 (0.465)	0.318 (0.322)
2014	1.018 (0.047)	0.609 (0.316)	0.237 (0.446)	0.315 (0.317)
2015	0.865 (0.056)	0.597 (0.318)	0.240 (0.455)	0.310 (0.308)
Jan - Jun 2016	0.757 (0.049)	0.586 (0.321)	0.237 (0.457)	0.304 (0.299)
Correlation matrix:				
P_{it}	1.000	0.105	0.045	0.007
C_{it}^{HHI}		1.000	-0.305	-0.675
MM_{it}^{I}			1.000	0.482
MM_{it}^{II}				1.000

The number of observations in the full period and the correlation matrix is 186,832. The price variable (P_{it}) is expressed in euros/liter, and calculated net of taxes (following information from the Spanish Tax Agency).

Table 2: Regression results

Dependent variable: lnP_{it}	OLS (a)	OLS (b)	OLS (c)	2SLS (a)	2SLS (b)	2SLS (c)
C_{it}						
C_{it}^{HHI}	0.715***	0.895***	1.414***	0.816***	1.027***	1.893***
	(0.063)	(0.063)	(0.145)	(0.071)	(0.070)	(0.200)
MM_{it}						
MM_{it}^{I}		0.496***			0.597***	
		(0.049)			(0.065)	
MM_{it}^{II}			0.846***			1.563***
			(0.212)			(0.274)
$C_{it}MM_{it}$						
$C_{it}^{HHI}MM_{it}^{I}$		-0.529***			-0.716***	
		(0.092)			(0.111)	
$C_{it}^{HHI}MM_{it}^{II}$			-0.047			-0.785**
			(0.246)			(0.300)
R^2	0.980	0.980	0.980	0.982	0.982	0.982
AIC	-974,432	-974,671	-974,890	-773,054	-773,224	-773,304
Total observations	186,832	186,832	186,832	149,873	149,873	149,873
Hansen J				6.015	8.858	12.203
				[0.198]	[0.546]	[0.272]
K-P rk Wald F				13×10^4	3.7×10^4	1×10^4
K-P rk LM				17.968	17.977	17.799
				[0.003]	[0.082]	[0.086]
Durbin-Wu-Hausman				24.16	30.66	44.26
				[0.000]	[0.000]	[0.000]

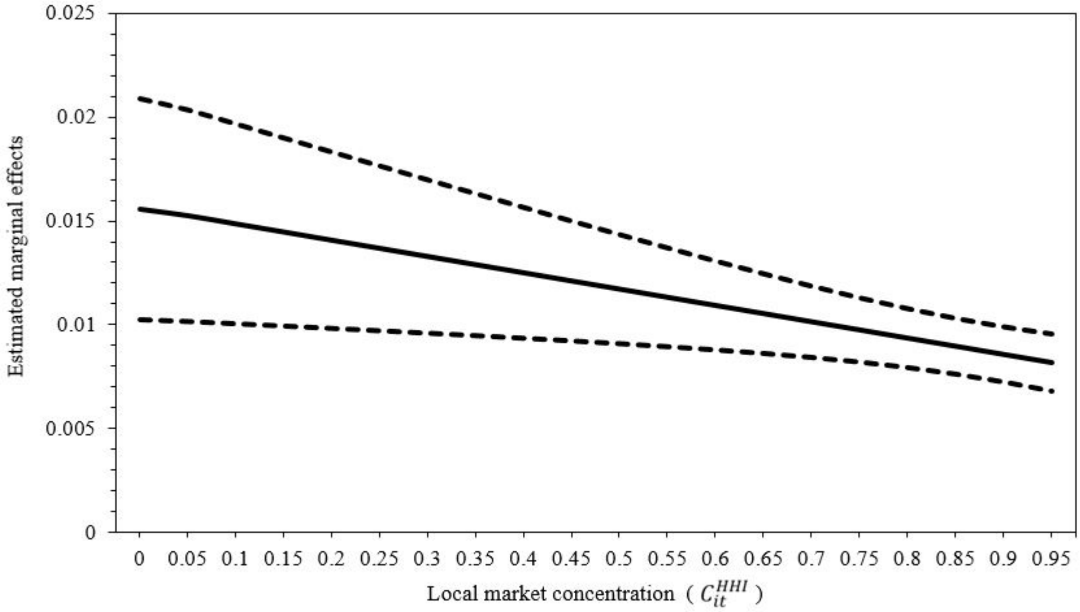
Regressions contain dummy variables for time, postcodes, and station brands. Driscoll-Kraay's standard errors are presented in parenthesis. *** and ** denote statistical significance at 1% and 5% levels, respectively. P-values are in brackets. Stock and Yogo critical values are used in Kleibergen-Paap (K-P) Wald F statistics. For the 2SLS estimator, the variables on the right-hand side of Equation (1) are instrumented by their own lagged values (up to four quarters) and the municipality-specific population. Coefficients and standard errors are multiplied by 10^2.

Table 3: Robustness check based on the 2SLS procedure

Dependent variable: lnP_{it}	(a)	(b)	(c)	(d)	(e)	(f)
C_{it}						
C_{it}^{HTI}	0.454***	2.390***				
	(0.046)	(0.272)				
C_{it}^{RI}			0.175***	1.455***		
			(0.042)	(0.218)		
C_{it}^{CCI}					0.783***	2.028***
					(0.077)	(0.241)
MM_{it}						
MM_{it}^{II}		3.399***		2.117***		2.267***
		(0.468)		(0.358)		(0.323)
$C_{it}MM_{it}^{II}$						
$C_{it}^{HTI}MM_{it}^{II}$		-3.682***				
		(0.468)				
$C_{it}^{RI}MM_{it}^{II}$				-2.466***		
				(0.357)		
$C_{it}^{CCI}MM_{it}^{II}$						-2.344***
						(0.340)
R^2	0.982	0.982	0.982	0.982	0.982	0.982
AIC	-772,523	-772,701	-772,408	-772,500	-772,954	-773,028
Total observations	149,873	149,873	149,873	149,873	149,873	149,873
Hansen J	5.429	14.302	7.313	14.221	6.865	12.958
	[0.246]	[0.160]	[0.120]	[0.163]	[0.143]	[0.226]
K-P rk Wald F	2.3×10^4	233.128	2.7×10^4	330.549	2.7×10^4	275.619
K-P rk LM	17.932	17.619	17.973	17.805	17.946	17.753
	[0.003]	[0.091]	[0.003]	[0.086]	[0.003]	[0.088]
Durbin-Wu-Hausman	17.88	19.29	6.77	5.88	12.25	44.26
	[0.000]	[0.000]	[0.009]	[0.015]	[0.000]	[0.000]

Regressions contain dummy variables for time, postcodes, and station brands. Driscoll-Kraay's standard errors are presented in parenthesis. *** denotes statistical significance at 1% level. P-values are in brackets. Stock and Yogo critical values are used in Kleibergen-Paap (K-P) Wald F statistics. The variables on the right-hand side of Equation (1) are instrumented by their own lagged values (up to four quarters) and the municipality-specific population. Coefficients and standard errors are multiplied by 10^2.

Figure 1. Marginal effects of MM_{it}^{II} on $ln(P_{it})$ from the 2SLS estimation



Note: Dashed lines represent 95% confidence intervals.

Appendix A. Multimarket contact measures

Our measures of multimarket contact MM^{I}_{it} and MM^{II}_{it} are inspired by those defined in Jans and Rosenbaum (1997) (pages 408-410) and Coccorese and Pellecchia (2013) (pages 194-195). That is, for each quarter, we begin by defining the matrix U, whose elements u_{ki} take the value 1 or 0, depending on whether or not firm $k = 1, \ldots, K$ owns/operates at least one selling point in the local market surrounding each station $i = 1, 2, ..., n$. Matrix U is then employed to build the contact matrix $A = UU'$, whose off-diagonal elements $a_{pq} = \sum_{i=1}^{n} u_{pi}u_{qi}$ quantify the number of markets where each pair of firms p and q meet, while their diagonal elements represent the number of markets where each firm operates.

Bearing in mind the previous definitions, the first measure of multimarket contact is constructed as follows for each quarter:

$$MM^{I}_{i} = \frac{\sum_{p=1}^{K-1}\sum_{q=p+1}^{K} a_{pq}u_{pi}u_{qi} - K_i(K_i-1)/2}{K_i(K_i-1)/2} \qquad \text{(Appendix .1)}$$

where the numerator measures the number of non-home market contacts for each pair of firms competing within the home market, while the denominator quantifies the total number of possible pairs of firms competing in such a home market.

The second measure of multimarket contact extends the first one by weighting the non-home contacts according to the firms' relative presences. To do so, we first construct the matrix B with the elements $b_{pqi} = \sum_{j=i}^{n} u_{pj}u_{qj}(\frac{n_{pj}}{N_j} + \frac{n_{qj}}{N_j})$, where n_{kj}/N_j is the number of stations associated to the firm k over the total number of stations operating in the local market j. The off-diagonal elements b_{pqi} then aggregate the relative presences for firms p and q across all non-home markets where they meet. Hence, for each quarter, we have:

$$MM^{II}_{i} = \frac{\sum_{p=1}^{K-1}\sum_{q=p+1}^{K} b_{pqi}u_{pi}u_{qi} - K_i(K_i-1)/2}{\sum_{p=1}^{K-1}\sum_{q=p+1}^{K} a_{pq}u_{pi}u_{qi} - K_i(K_i-1)/2} \qquad \text{(Appendix .2)}$$

which capture the relative presences averaged over the total number of non-home market contacts.