Universität St.Gallen

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June 2020 Discussion Paper no. 2020-09

School of Economics and Political Science,
Department of Economics

University of St.Gallen

Editor:	Vanessa Pischulti University of St.Gallen School of Economics and Political Science Department of Economics Müller-Friedberg-Strasse 6/8 CH-9000 St.Gallen Phone +41 71 224 23 07 Email seps@unisg.ch
Publisher:	School of Economics and Political Science Department of Economics University of St.Gallen Müller-Friedberg-Strasse 6/8 CH-9000 St. Gallen Phone +41 71 224 23 07
Electronic Publication:	http://www.seps.unisg.ch

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[1] We thank participants in seminars at CREST, University of Gothenburg, University of Mannheim, IFN (Stockholm), Swedish Competition Authority (Stockholm), Swedish Energy Economics Conference (Lulea), EEA (Geneva) and EC 2 (Edinburgh) for useful comments. Special thanks to Pierre Fleckinger for many insightful discussions. Petyo Bonev and Magnus Söderberg also acknowledge the generous financial support from the Swedish Competition Authority, grant number 372/2016.

Abstract

We examine a novel regulatory mechanism, in which otherwise unrelated monopolists are in implicit price competition. The mechanism resembles the classical yardstick competition, however the reference standard is set not by the regulator but by the customers. They demand fair prices and infer about the price fairness through a spatial price comparison. We test this mechanism using a unique dataset on unregulated District Heating natural monopolists in Sweden. To account for the endogeneity of reference prices, we use a novel instrument that utilizes a shock on the demand side triggered by an unanticipated policy reform. We find a large effect of reference prices on prices, which indicates that the implicit yardstick competition has a considerable disciplining effect on monopoly prices.

Keywords

Monopoly, price setting, spatial interaction, natural experiment, yardstick competition

JEL Classification

L12, L43

1 Introduction

This article examines a novel regulatory mechanism in a setting with multiple local natural monopolists. If these monopolists are subject to no, or weak, price regulation, customers form opinions about prices by comparing the prices set by adjacent monopolies. This 'implicit yardstick competition' has the potential to discipline firms that set prices that deviate from their neighbours' prices. In this paper we evaluate the effectiveness of the implicit yardstick competition in the context of the district heating (DH) sector in Sweden. The Swedish DH industry comprises a unique setup. It consists of 262 local municipality-level monopolies. Some of the monopolists are public, but a sizable share is privately owned. Contrary to common practice in such natural monopoly markets, prices are not regulated and there is no franchising.

Shleifer (1985) coined the term "yardstick competition" to describe a regulatory approach whereby a regulator sets the price of a regulated firm as a function of the costs of other comparable monopolists. This approach has been implemented in many regulated sectors (e.g. water supply, hospital healthcare).

We argue that in the Swedish DH case, the unregulated monopolists are disciplined by benchmarking performed by the customers. This claim originates in behavioral economics. An important theoretical and empirical result in this literature is that consumers rely on standards to judge whether a price is 'fair' and fairness standards influence their behavior (for example, Kahneman et al. (1986) and Rotemberg (2011)). The production cost provides an obvious standard, but consumers cannot easily infer its value in industries with complex production processes. In such cases, they use other references. A natural choice for customers is to compare the price with the prices for the same good set by other firms, Kahneman et al. (1986), Rotemberg (2005) and Rotemberg (2011). In the case of the Swedish DH industry, this is the role played by the neighbouring prices.

Using other firms' prices as a fairness standard creates a benchmarking system that disciplines firms. A decrease in the reference price reduces customers' utility. As a result, the firm sets a lower optimum (monopolist) price. This mechanism creates negative externalities between otherwise unrelated monopolists and it thus triggers an implicit price competition.

The major contribution of our paper is to empirically test for the existence of these negative externalities. This is a challenging task since a spatial correlation in prices can be also triggered by unobserved correlated factors of supply and demand. To deal with the endogeneity of reference prices, we develop an elaborate empirical strategy that utilizes three different sources of identification. In our main approach, we use a novel natural experiment as an instrument for the endogenous prices. The experiment is triggered by a demand-side shock caused by an unanticipated policy reform. Our second strategy is based on a parametric assumption about the joint distribution of the disturbances. Our third strategy, Indirect Inference (II), is based on simulating an auxiliary model, that is assumed to be encompassed by the true model. These three strategies - a moment condition, a parametric assumption about the unobservables and the II assumption - cover the vast majority of empirical strategies used in spatial econometrics. We find positive spatial positive price spillovers consistently in all three approaches. These spillovers are direct effects of DH prices in neighboring markets on the DH price in a market, and thus cannot be explained by correlation in confounding factors. Our estimates are sizeable, which implies that the disciplinary effect of the implicit yardstick competition is economic relevant.

An additional testable implication of the implicit yardstick mechanism is that publicly owned monopolists are less responsive to changes in reference prices. Our results confirm this hypothesis.

The paper is structured as follows. In section 2, we describe the institutional setup and the data. In section 3, we describe the mechanism to be tested and the empirical strategy. In section 4, we present our empirical results. In section 5, we conclude.

2 Institutional setup and data

2.1 Institutional setup

Currently, there is one DH firm in the main locality in 262 of the 290 municipalities.[1] All DH utilities are vertically integrated, i.e. production and distribution are owned by the same firm. With the exception of two large firms (E.ON and Fortum) and a few smaller collaborators that own networks in several municipalities, each firm/market is economically and legally independent from all other firms/markets.[2] The high fixed distribution costs and the fact that customers can only purchase DH from the firm in the municipality where they resides imply that each utility is a local natural monopoly.

Furthermore, DH technology is only viable in densely populated urban areas. Customers in cities typically have two possible sources of heat: DH and electricity-based technologies, primarily in the form of heat pumps.[3] For customers connected to the DH network, DH is the cheapest source of heat compared to electricity with a ratio of variable costs around 0.5, EMI (2012). Due to the geographical restriction of DH and the high switching cost[4], these customers are locked in to their providers. This lock-in effect and the fact that heating is a basic need in Sweden lead to a demand elasticity close to zero, see Brännlund et al. (2007). In contrast, the change in demand that results from attracting new customers is elastic and sizable, Biggar et al. (2018).

The DH market opened for private investors in 1996. The DH prices are set independently by each firm, and since 1996, prices are not subject to any periodic sector-specific review by a regulatory agency. This is in stark contrast to how electricity prices are set: the retail price is determined on a competitive market and the transportation prices

[1] An extended description of the institutional setup and the data can be found in our companion paper Bonev et al. (2018).

[2] Firms that are active in several municipalities are excluded from the empirical analysis.

[3] Natural gas plays a negligible role in Sweden, and oil was practically phased out during the 1980s and 1990s.

[4] A one-time connection fee is paid at the time when the residence connects to the network. This fee is high and can amount to ten times the total annual consumption cost.

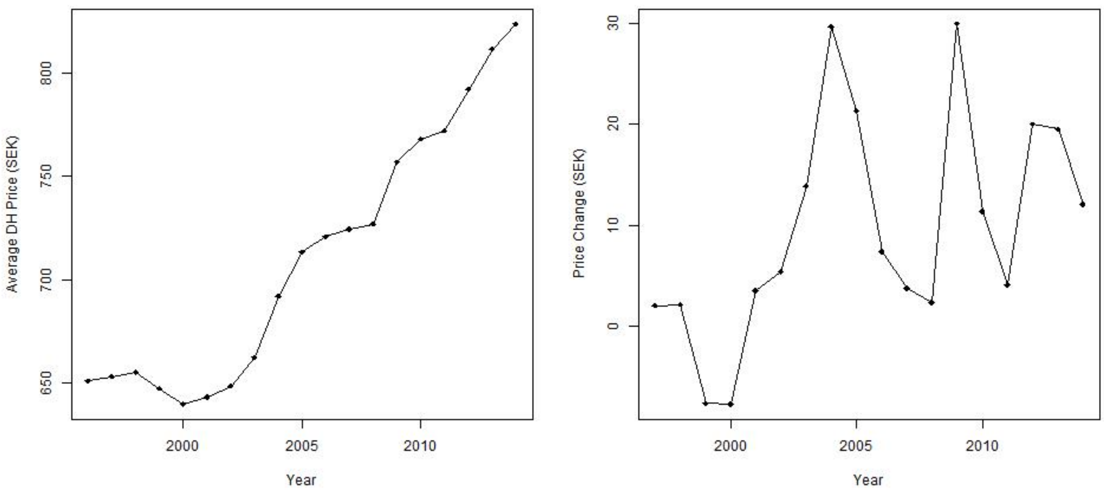
(transmission and local distribution) are regulated by the Swedish Energy Markets Inspectorate through ex ante revenue caps.

2.2 Data and descriptive statistics

Our dataset contains annual information on each local DH market. Information about prices was gathered from the Nils Holgersson annual price survey (NHS), which reports municipal specific list prices for a representative customer.[5] Between September and November each year, the firms announce the price that will be charged in the next year from the 1st of January. Figure 1a shows the average price for DH for each year between

Figure 1: Average prices and price changes.

(a) Average prices, 1996-2014 (b) Changes in average prices



1996 and 2014, and 1b shows the changes in price over the same time period.

Additional demand and supply characteristics were added from other sources. On the

[5] This annual survey is run by several of the largest organizations with interest in the Swedish property markets, specifically the Swedish Union of Tenants, HSB Riksfoerbund (Sweden's largest housing cooperation), Riksbyggen (an organization owned by the building unions, local housing associations and by other national co-operative associations), SABO (the Swedish Association of Public Housing) and the Swedish Property Federation.

supply side, we gathered data on the different fuel types and quantities used to produce heat by each firm. The types and amounts of fuel used affect firms' cost levels. This information was collected from the Energy Markets Inspectorate and directly from the DH firms. We also gathered data on labor cost. On the demand side, our dataset contains municipal-specific information on the average income, the share of the population above the age of 65, the total number of inhabitants, and the share of detached dwellings (i.e. single family houses). These variables were collected from Statistics Sweden and from the municipalities. Additional covariates such as the electricity tax (a measure of the price of the substitute) and weather (number of heating degree days and amount of precipitation) were also gathered. Since they have no (electricity tax) or only very limited (weather) cross-sectional variation, including them into the analysis made no difference. In Table 1, we show descriptive statistics for the period 2008-2009. The

Table 1: Summary statistics of the sample(2008-2009)

Variable	Min	1st Qu	Median	Mean	3rd Qu	Max
Price 2009	423	678.3	743.5	728.7	783.6	912.1
Price 2010	437.8	711.5	771.7	761.2	818.1	964.9
Rel. Δ Price	-0.10	0.024	0.042	0.045	0.059	0.28
Rel. Δ Population	-0.014	0.0006	0.002	0.003	0.005	0.023
Rel. Δ Labor cost	0.011	0.014	0.015	0.014	0.016	0.017
Rel. Δ Age > 65	-0.054	-0.019	-0.009	-0.011	-0.003	0.016
Z_{i2009}	0.025	0.21	0.62	0.60	0.71	0.84

Note: Rows 3-8 contain statistics for the relative change (Rel. Δ) 2008-09 of a variable. Z_{i2009} is the local share of detached houses in 2009. Summary based on 225 observations.

choice for this time span is motivated in the identification section below. The table contains descriptive statistics for the local price levels in 2008 and 2009, as well as for the relative changes of all variables in that time period. The relative change variable is defined as (Variable 2009 - Variable 2008)/Variable 2008. The relative changes of the observed covariates (other than prices) between 2008 and 2009 have only small variation across units. The share of population of age over 65, for example, has a first quartile equal to a 1% decrease and a 3rd quartile equal to a 0.3% decrease. This finding is important for the interpretation of our results in section 4. The relative price change, on the other hand, exhibits substantial variation, with largest price change being 28% and the lowest being negative.

The last variable in table 1, denoted by Z_{it}, is explained in the next section.

The substitute. The only substitute of district heating is electricity-based heating. The electricity price paid by end-consumers consists of three parts - an electricity tax the retail price and the local distribution price - where only the first component is observable. The increase in electricity tax in the period 2010-2013 has practically no cross-sectional variation.[6] The retail price is determined on a highly competitive international market (the so called Nord Pool electricity retail market) and has generally no local variation, see Botterud et al. (2010) for a detailed description of the Nord Pool market. The Swedish Energy Markets Inspectorate records prices for every regulated local distribution firm. According to these records, the distribution price has been almost constant during the relevant sample period. As a result, it can be plausibly argued that the price of the substitute can be modeled as an individual time-constant fixed effect.

Publicly own DH firms. As of 2008, About 60% of all the DH utilities are owned by the municipalities and the remaining 40% is owned by private and foreign investors. Figure 1a provides an overview of Swedish DH networks in the 290 municipalities (light gray represents public ownership and dark gray private ownership. White municipalities either did not have a DH network in 2008 or the information about their ownership form is incomplete. Figure 1b displays the annual absolute change in prices for public (thick line) and private municipalities (dashed line) in the period 2000 - 2014. In the main period of evaluation 2008-2009, average prices were almost identical.

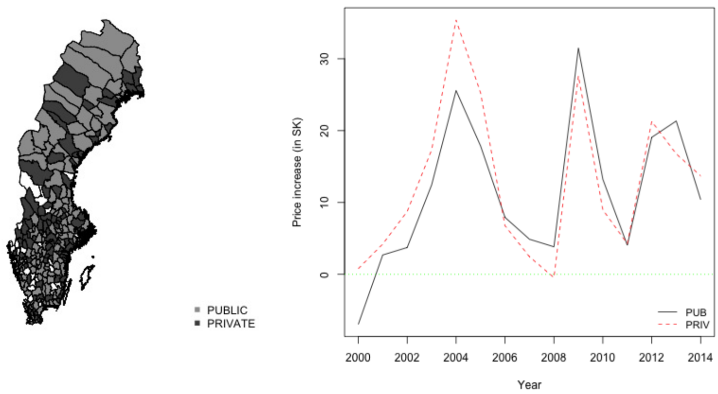
Finally, table 2 contains a comparison of observed characteristics of privately owned

[6] Descriptive statistics is available upon request.

Figure 2: Left: private (dark gray) vs. public (light gray) municipalities. Right: absolute price increase.

(a) Private vs. public

(b) Price changes private vs. public



and public firms (and municipalities) in 2009. The last column contains the p-values of

Table 2: Comparison of average observed characteristics of privately and publicly owned DH firms

Variable	Public	Private	p-value
# Households	17178.9	16127.5	0.82
Age DH	25.45	21.15	0.004
Av. Income	248.55	243.29	0.21
Age < 15	0.162	0.161	0.73
Electricity	274.96	284.77	0.03
Weather	4412.68	4243.81	0.05

a t-test of equality of means. Municipalities with a public owner of the DH firm have a higher price for the substitute electricity. Also, these regions are colder. On the other hand, demand properties such as local average income and age distribution do not differ significantly.

The local flow of information. The following brief exposition will be used later to motivate our spatial analysis. First, local newspapers are the major source of information on DH prices for customers. These newspapers typically cover 2-4 municipalities.[7] Second, a big share of the working population commutes to one of the neighboring municipalities.[8] In addition, in Sweden, there are 72 local labor markets (data 2014) defined on an administrative basis, which corresponds to 4.03 municipalities per labor market on average - close to the average number of direct neighbors of a municipality.[9]

3 Mechanism, notation and empirical strategy

Consider a consumer who can purchase DH in market i at price p_i. Drawing on findings in the behavioral literature, we assume that customers demand a fair price, Kahneman et al. (1986), and expect firms to be benevolent, Rotemberg (2011). In particular, as in Rotemberg (2005) and Rotemberg (2011), assume that the utility u of a customer depends not only on the physical properties w of the product, but also on a parameter α that captures the customer's assessment of the firm's altruism:[10]

$$u = u(q; \alpha), \tag{1}$$

where q is the quantity DH consumed by the customer. We normalize the relationship between u and α to be increasing,

$$\partial u / \partial \alpha > 0, \tag{2}$$

that is, a higher α reflects a higher degree of belief in the altruism of the firm. Following Kahneman et al. (1986) and Rotemberg (2011), we assume that customers rely on fairness

[7] A report on the Swedish media landscape in Sweden can be found on the webpage of the European Journalism Centre, http://ejc.net.

[8] See e.g. http://www.grs.scb.se.

[9] This information is obtained from the Statistics Sweden, which is the government statistics institute in Sweden: http://www.scb.se/en/

[10] In our setup, the product is a homogeneous good. The dependence on the physical properties will be therefore omitted.

standards when they assess the fairness of the price. The production cost provides an obvious standard, but consumers cannot easily infer its value in industries with complex production processes. In such cases, customers use other references, in particular, the prices set by firms producing a similar type of goods or services, Kahneman et al. (1986). If p_i is too high compared to the average price in close markets, the price is deemed unfair. To capture this relation, let α be a function of the difference between the price charged p_i and the average price in other markets p_{-i}:

$$\alpha = \alpha(d_i) \qquad \text{with} \frac{\partial \alpha}{\partial d} \leq 0, \tag{3}$$

where $d_i = p_i - p_{-i}$. In practice, the consumer exhibits high search cost which is proportional to distance, see the end of section (2.2) ("the local flow of information"). Thus, it is reasonable to assume that customers compare their price only with prices in close markets. In our companion paper Bonev et al. (2018), we provide comprehensive evidence for for the validity of (3). In particular, we proxy α with the number of customer complaints to a national DH board, and show that d_i has a causal effect on α_i. In the following, p_{-i} will refer to the average price in close markets. Specifically, we define

$$p_{-i} = \sum_{j \neq i} w_{ij} p_j,$$

where w_{ij} are non-negative spatial weights that sum up to 1. If market j is considered as close to market i, then w_{ij} is strictly positive.The precise meaning of "close" is defined below. Following an established convention in spatial econometrics, w_{ii} is set to 0 for all i.

(1), (2) (3) imply that an exogenous decrease in the average price p_{-i} leads to a decrease in p_i:

$$\frac{\partial p_i}{\partial p_{-i}} \geq 0. \tag{4}$$

It follows that the reference-prices mechanism creates a relation between otherwise unrelated monopolists. These monopolists are in an implicit price competition: one firm lowering its price might trigger a decrease in prices in other markets. As a result, an a non-cooperative equilibrium, the prices set are lower then the individual optimal monopolist prices that would be set in absence of other markets.

In the following, we focus on testing the relationship (4). Importantly, the correlation generated by the references-price mechanism reflects a causal effect of p_{-i} on p_i. It is thus different than the correlation in prices caused by correlated factors of supply and demand. It is a major challenge of identification to disentangle these two types of correlation. In particular, potentially unobserved correlated factors of demand and supply render the spatial price lag p_{-i} potentially endogenous. A second source of endogeneity is that prices are determined simultaneously. p_{-i} on p_i are thus in a reverse causality relationship.

We use three different identification and estimation strategies. The first one relies on an exclusion restriction and is presented in great detail below. The second one relies on a specification of the distribution of the error term as a multivariate normal distribution and uses a maximum likelihood estimator. This is a standard technique in spatial econometrics and is not discussed separately. The third one is the so-called indirect inference (II) approach. We briefly sketch the idea behind II in the appendix.

3.1 Identification with an instrument

To instrument for p_{-i}, we utilize a natural experiment triggered by a policy shock. On December the 5th, 2008, the Swedish government announced that households in detached houses would be subject to an optional subsidy of 50,000 SEK (equivalent to about $4,800$ € in December 2008) to increase the energy efficiency of their homes. The subsidy could be used for any type of measures that would improve the energy efficiency of the dwelling, including the connection to district heating.

The time structure of this policy, depicted in figure 3, was exploited to construct our instrument. The dashed vertical line denotes the date of the announcement of the policy on the 5th of December 2008. The official start was only four days later as the subsidy could be claimed from the 9th of December. The last possible day for claiming the subsidy was 30th of June 2009, or about 7 months after its implementation. As explained above, DH prices are adjusted only once per year. More precisely, firm i commits to price p_{it} between September and November in year $t-1$.[11] Thus, the choice of the 2009 price took place before the introduction of the subsidy, and the 2010 price was chosen after the last day that the subsidy could be claimed. The two price setting periods are depicted as shaded rectangles in figure 3.

Prior to the official announcement of the subsidy, there was limited debate about it. The proposal regarding the subsidy was first debated in the Swedish national parliament on the 8th of October, two months before its announcement and implementation.[12]. We have no data on the precise dates that each firm announced their prices for 2009, and some firms might have done so after the 8th of October 2008. However, the legislative discussion left much room for uncertainty as there were no clear outcome and announcement regarding the precise content and timeline. It was also not clear for which components the subsidy could be used.

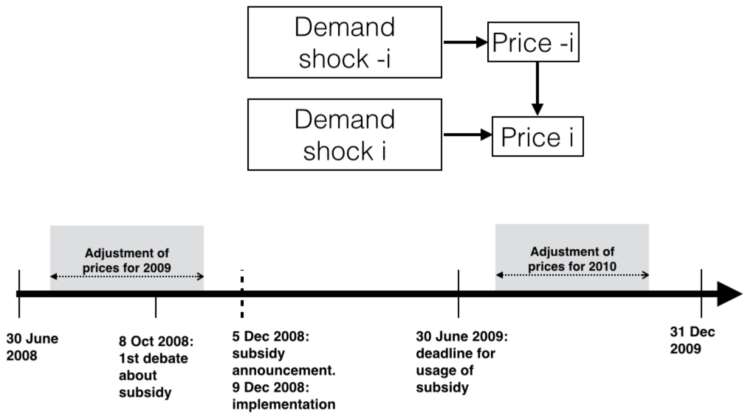
One effect of this subsidy was to make connecting to DH cheaper in 2009, which is likely to have attracted new consumers. Since the subsidy could only be taken until June 2009, this change in the stock of customers must have occurred prior to adjustment of prices for 2010. As a result, the subsidy is likely to have induced a change in the 2010 prices. We exploit variation in the induced change in customer stock across markets as an instrument for p_{-i}. In particular, as depicted in figure 3, a shock in the demand stock in markets $-i$, neighboring with market i, is likely to have changed the price in

[11] See http://www.sevab.com/Privat/Fjarrvarme/Priser/ for more information and examples (in Swedish).

[12] The link to the official report of the government about this debate is https://www.riksdagen.se/sv/Dokument-Lagar/Forslag/Motioner/Inforande-av-ROT-avdrag_GW02Sk379/?text=true .

markets $-i$, but it had no direct effect on the price in market i (the logic behind the exclusion restriction is discussed in detail below). This identification strategy is related to the one used by Lyytikäinen (2012), who uses local variation in tax increases induced by a national policy reform to identify the spillover effect of local tax levels in a tax competition setting. Our strategy is also related in spirit to the natural experiments used to test the Permanent Income Hypothesis, as the majority of these settings rely on an unanticipated income shift, see Fuchs-Schuendeln and Hassan (2015) for a detailed overview and discussion.

Figure 3: Timing and causal structure of the change in customer stock.



Unfortunately, we do not observe changes in the stock of customers triggered by the subsidy as we have no information on individual households' decisions to claim the subsidy. Instead, as a proxy we use the following variable

$$Z_i := \frac{\text{Number of households in detached houses in municipality i}}{\text{Total number of households in municipality i}}. \tag{5}$$

The interpretation of Z_i is as follows. The nominator in (5) measures the maximum number of households that would have been eligible for the subsidy if no detached households had yet connected to DH. This is a (fictive) upper bound for the number of new customers that could have been gained by a DH firm due to the subsidy. The denominator in (5) measures the total size of the demand of the DH firm in the fictive case where all households are customers. Thus, the variable Z_i provides a measure of the maximum potential change in customer stock in municipality i due to the subsidy. It is thus similar to an Intention-to-Treat variable that results from the treatment assignment in a randomized experiment, see Heckman et al. (1999).

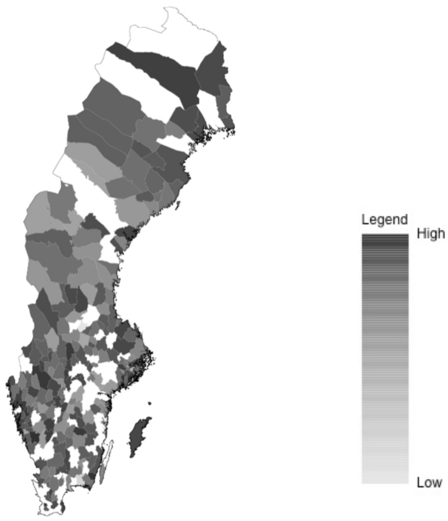
In a first-difference regression, we use $Z_{-i2009} = Z_{-i2009} - 0$ as an instrument for the endogenous price difference $p_{-i2009} - p_{-i2008}$, where Z_{-it} is defined as the average value of Z at time t in markets close to market i. In particular, $Z_{-it} = \sum_{j \neq i} w_{ij} Z_{jt}$. The spatial weights are identical to those in the definition of p_{-it}. Thus, implicitly, each change of the demand stock in a neighboring market instruments for the price in that market. Z_{it} exhibits rich variation over its support $[0, 1]$. Roughly 90% of all values of $Z_{i,2009}$ lie between 0.2 and 0.8. The minimum of Z_i is 0.025 and the maximum is 0.84, with an average of 0.599 and 3rd quartile of 0.71. The correlation of Z_{-i2009} and Z_{i2009} amounts to 0.13. The p-value of the Moran's I test is 0.11, so the test fails to reject the null hypothesis of no spatial correlation at the 10% level. In addition, a graphical inspection of the spatial distribution of Z_{i2009} reveals no visible patterns of dependence, see figure 4.

We now discuss in detail the properties of Z_{-it} that qualify it as a valid instrument.

Exclusion restriction. The percentage of detached households in market $-i$ is a valid exclusion restriction since it has no direct effect on the price in market i due to the economic and legal independence of the markets as discussed in section 2.1. The exclusion restriction is depicted in figure 3 as the lack of a direct link between Z_{-it} and p_{it}.

No anticipation. Neither customers nor DH firms anticipated the policy shock. The time span from the first debate on the 8th of October to the actual decision to

Figure 4: Spatial distribution of the instrument. Darker color indicates values closer to one and white regions indicating missing prices.



introduce the subsidy was only 8 weeks and the discussion on the 8th of October left room for uncertainty on the design of the subsidy. In addition, the time span between the official decision on the 5th of December and the actual implementation was only 3 days. No anticipation of the reform is an important characteristic, as it precludes forward-looking unobserved behavior that could potentially influence the prices. One example for such behaviour would be firms adjusting their 2009 prices in anticipation of a future shift in demand.

Exogeneity of the instrument. Next, we justify the exogeneity assumption $Z_{-it} \perp\!\!\!\perp \varepsilon_{it}$. We rely on four arguments. First, it can be plausibly claimed that the instrument is not related to unobserved variable costs. In particular, labor costs and fuel costs are the main variable costs of district heating, amounting to more than 90% of the total variable cost.[13] While labor cost is observed directly and appears to be uncorrelated with Z_{-i2009}, the fuel cost has virtually no local variation. In particular, Sweden

[13] See e.g. Difs and Trygg (2009) and Sjödin and Henning (2004) for how to calculate the DH marginal cost and the relation between marginal cost and price.

is divided into 3 large "fuel regions". Within each of those regions, markets are exposed to the same price for most fuels, including bio-fuel (which is the major fuel type used by DH firms). Thus, variation in fuel cost is not related to local market characteristics.[14]

Second, we regress the instrument on Z_{-i2009} on *observed* factors X_{it} of the price in market i. The results are presented in Table 3. Although the estimated effect of

Table 3: Regression of instrument on observed covariates

	Coef.	Std. Err.	t	p-value
Intercept	0.57^{**}	0.024	22.98	2e-16
Population	-0.00001^{**}	4.82e-06	-2.07	0.04
Age> 65	5.04e-06	2.32e-05	0.217	0.82
Av. Income	0.004	0.003	1.32	0.19
Num of obs:	228			

Note: regression of the instrument Z_{-i} on observed covariates in market i. Column 2 displays the estimated coefficients and the last column the p-values.

the number of households is significant, the magnitude of the coefficient is economically irrelevant. All other observed covariates have insignificant estimated coefficients. By way of analogy, these insignificant results provide indirect evidence that the instrument is not related to the *unobserved* factors ε_{it} of the price in market i.

This interpretation is reinforced by our third argument, namely that Z_{i2009} itself is not significantly spatially correlated. Its behavior, as depicted above, resembles an idiosyncratic shock. This indirect argument draws on the following observations. Observed factors of supply and demand do exhibit high spatial correlation (e.g. fuel cost, electricity tax and weather conditions, almost 100%, labor costs over 80%). Again by way of analogy, one can assume that a similar correlation holds for unobserved factors. Thus, if the instrument was related to these unobserved factors, one would expect to find spatial correlation in its distribution as well.

[14] See the website of the Forest Statistics Yearbook, where prices for most fuel types are reported: http://www.skogsstyrelsen.se/en/AUTHORITY/Statistics/Statistical-Yearbook-/Statistical-Yearbooks-of-Forestry/

Finally, as documented in section 2.2, the price of the only substitute (electricity) has no time and regional variation for the period of consideration. As a result, there was no unobserved variation in competition intensity that could potentially be related to Z_{-it}.

4 Estimation

We estimate the model

$$p_{it+1} = \beta_0 + \rho \sum_{j \neq i} w_{ij} p_{jt+1} + X_{it}\beta + \sum_{l=1}^{T-1} \delta_l T_l + \theta_i + \varepsilon_{it}, \tag{6}$$

where X_{it} is a $1 \times k$-dimensional random vector of observed covariates, T_l are time dummies with $T_l = 1$ if $l = t$ and 0 otherwise, θ_i are municipality fixed effects, ρ, β_0, $\beta = (\beta_1, \ldots, \beta_k)'$ and $\delta_1, \ldots, \delta_{T-1}$ are unknown coefficients and ε_{it} is the time-varying idiosyncratic error of the regression model. The main parameter of interest is ρ. It can be interpreted the first derivative of p_{it} with respect to the average of neighboring prices p_{-it}, $\rho = \partial p_{it} / \partial p_{-it}$. It reflects the strength of the spatial spillover of neighboring prices on the price in market i.

To utilize the time structure of the policy reform described in the last section, we take the first-difference of two consecutive periods and estimate

$$p_{i2010} - p_{i2009} = \beta_0 + \rho \sum_{j \neq i} w_{ij} (p_{j2010} - p_{j2009}) + (X_{i2009} - X_{i2008})\beta + (\varepsilon_{i2009} - \varepsilon_{i2008}). \tag{7}$$

Spatial lags of covariates are not included due to the economic and legal independence of the markets.

Table 4 presents the results of six different estimation procedures. The first four estimates are Instrumental Variable estimates with an instrument Z_{-it}. Estimation (5) is obtained using a Maximum Likelihood Estimation (MLE) procedure, and estimation

(6) is obtained using Indirect Inference (II) estimation procedure, see appendix A for a brief description and a reference. The three estimation strategies rely on three different identifying assumptions: (i) valid instrument, (ii) correctly specified distribution of the idiosyncratic term, and (iii) homogeneous iid disturbances. Within the IV estimations, we also vary the choice of the weights. Specifications (1) - (3) (and also (5) and (6)) assume weight $w_{i,j} = 1$ if market j shares a border with market i and 0 otherwise. Specification (4) assumes weights proportional to 1/distance if market j for bordering markets and 0 otherwise. Alternative choices of the weights yield results very similar to the results in table (4) and are therefore not presented. Furthermore, the IV specifications differ w.r.t. the underlying assumptions about the error term. Specifications (1) and (4) ares estimated with the spatial TSLS estimator developed in Kelejian and Prucha (1998). The standard errors are heteroskedasticity-robust, and the disturbances ε_{it} are assumed to be independent. In specification (2), we allow for spatial dependence in the disturbances of the model. This dependence could arise due to spatially correlated costs or demand characteristics. The disturbances are modeled according to the standard spatial autoregressive model with autoregressive disturbances of order (1, 1) (SARAR(1, 1)), see e.g. Anselin and Florax (1995),

$$\varepsilon_t = \eta M \varepsilon_t + \xi_t, \tag{8}$$

where M is a $n \times n$ spatial weights matrix and $\xi_t = (\xi_{1,t}, \ldots, \xi_{n,t})$ is a vector of independent innovations with variances $\sigma_1, \ldots, \sigma_n$. The parameter η represents the coefficient of the spatial lag of the errors. Specification (3) assumes the same error structure as in specification (2) but adds further observed covariates.

All 6 regressions produce a positive estimate of ρ, and the first (5) estimates are significant.[15] The IV estimates vary between 0.55 and 0.77, thus all lying within the

[15] The II method does not provide a way to compute the standard errors.

Table 4: Empirical results.

	IV Estimates				MLE	II
	(1)	(2)	(3)	(4)	(5)	(6)
ρ	0.62**	0.77***	0.55**	0.47**	0.21***	0.37*
	(0.32)	(0.26)	(0.33)	(0.27)	(0.079)	(0.209)
Population			0.02			
			(0.03)			
Age > 65			-0.87			
			(0.58)			
Labour Cost			0.002			
			(0.004)			
Intercept	9.21	6.19	14.36	8.69	19.90***	15.76
	(11.16)	(6.43)	(13.50)	(11.93)	(6.11)	
η		0.40*				
		(0.21)				
I-Stage F-stat.	15.83	14.79	21.29	13.038		
Num of obs.:	229					

Price change is regressed on "average" DH price in neighboring markets and observed covariates. Specifications (1) -(4) are IV estimates with Z_{-it} as an instrument for p_{-it}. Specification (5) presents MLE estimates and (6) presents II estimates. Specifications (1)-(3), (5) and (6) use weights 1 for neighbors with shared border and zero otherwise (row standardized). Specification (4) assumes weights proportional to 1/distance. Specification (1) assumes independent disturbances, and the standard errors are heteroskedasticity robust. Specification (2) is a SARAR(1,1) model. Specification (3) includes observed covariates. * denotes $p < 0.1$, ** denotes $p < 0.05$, *** denotes $p < 0.01$. p-values are calculated for one-sided tests.

Table 5: First stage results corresponding to the IV estimation in Table 4.

	(1)	(2)	(3)	(4)
Z_{it}	33.12***	35.30***	30.46***	33.12***
	(7.395)	(8.16)	(12.45)	(8.45)
Population			0.09	
			(0.56)	
Age > 65			3.67	
			(14.56)	
Labour Cost			0.0003	
			(0.0012)	
Num of obs.:	229			
F-stat.	15.83	14.79	21.29	13.038

Note: Specifications (1)-(4) display the first stage estimates to specification (1)-(4) from Table 4, respectively. * denotes $p < 0.1$, ** denotes $p < 0.05$, *** denotes $p < 0.01$.

90 % confidence interval around 0.62 (the intermediate estimate). Since we instrument for p_{-it}, $\hat{\rho}$ can be interpreted as a causal effect of the weighted average price p_{-i} in neighboring markets on the price in market i. On the basis of these results, a unit increase in p_{-i} induces a change between 0.55 and 0.77 units of the price of firm i, which implies that the spatial spillover is economically strong. The MLE and II estimates are of smaller magnitude, with the II estimate closest to the lowest IV estimate. The estimates of the effect of other observed covariates in specification (3) are not significant. The finding is not surprising given the lack of variation of these covariates between 2008 and 2009 as documented in Table 1.

Further, the instrument is reasonably strong in all four regressions with the Kleibergen-Paap F statistics being between 13.038 and 21.29. The first stage results are summarized in Table 5. As an example, in the first stage of specification (1), the estimated coefficient of Z_{-it} is positive and has the value 33.12. Under the assumption that this coefficient has a causal meaning, a unit increase in the change of customer stock leads to a price increase of around 30 SEK. This corresponds to a 4% increase in the average price of DH in 2009. We remind that the average Z is equal to 0.6, which corresponds roughly to a 2.5% increase in the price change.

Thus, we obtain positive and significant estimates of the spatial spillover effect ρ under (1) different identifying assumptions, (2) different spatial weights, (3) different assumptions about the error term and (4) different sets of explanatory variables.

Estimation of spillover effects on the subsample of public firms. As another robustness check, we re-estimate regression (7) with a restricted sample. In particular, we include only those observations, whose ownership form is a public one. Note that the restriction affects solely which observations are included into the sample, not the values of the spacial lag and the instrument. In particular, for a given observation in the restricted sample, Z_{-i} and p_{-i} are the same as the corresponding values for that observation when the full sample is used.

The intuition behind this exercise is the following. If $\hat{\rho}$ reflects a causal price spillover effect, then it is plausible to expect that this effect should be of smaller magnitude for the restricted sample. First, customers in Sweden tend to grant higher level of trust to public DH monopolies (e.g. Palm and Magnusson (2009) find that price increases implemented by municipality-owned DH utilities are treated substantially less harsh in the media than when implemented by alternative owners). Second, publicly firms are likely to be less susceptible to market forces than privately owned.

Table 6 provides an overview of estimation results. Each line represents one specification. (1) corresponds to specification (1) in table 4, that is, a spatial TSLS. (2) correponds to specification (2) in table 4 and (3) - (5) to (4) - (6), respectively. The estimates obtained with spatial TSLS, (1) and (3), and with spatial GMM (2), are not significantly different from zero. The Kleibergen-Paap F statistics for these specifications is 10.68 and 11.05, so the instrument is above the rule-of-thumb level of 10. The ML (specification (4)) estimate is significant and positive, but of smaller magnitude than the on the full sample. Finally, the II estimate is also insignificant. Overall, these estimates are in line with our predictions.

Table 6: Results with the restricted sample (only public firms).

Specification	estimate	p-value	F stat
(1)	-0.19	0.672	10.68
(2)	0.12	0.41	10.68
(3)	-0.23	0.71	11.05
(4)	0.18**	0.012	
(5)	0.11	0.4	
Num of obs:	134		

Note: Specification (1) corresponds to specification (1) in table 4, specification (2) to specification (2), specifications (3)-(5) to specifications (4)-(6) . * denotes $p < 0.1$, ** denotes $p < 0.05$, *** denotes $p < 0.01$. p-values are calculated for one-sided tests.

5 Conclusion

In this paper, we present a mechanism that we refer to as implicit yardstick competition. Unrelated monopolies that sell a homogeneous good are related to each other through a price benchmarking system. In particular, customers demand a fair price and to infer about the fairness of the price, they compare its own price with prices charged in other, neighboring, markets. This referencing mechanism triggers negative externalities on the firms' prices and disciplines their pricing behavior.

We test this mechanism using a unique dataset on Swedish DH firms. All three empirical strategies reveal that the negative externalities are economically significant, implying that their disciplining power is substantial.

A Indirect Inference Approach

The II approach implemented here is based on the theoretical results of Kyriacou et al. (2017). This is a simulation method that can be described in the following way. Suppose that the "true" parameter ρ_0 lies in a closed subset Λ of $(-1, 1)$. For any element ρ of Λ, generate K datasets $y_1(\rho), \dots, y_K(\rho)$, each of them following the model .[16] For each data

[16] The error term is generated from the normal distribution.

set $y_k(\rho)$, calculate the OLS estimator of ρ_0, $\hat{\rho}_k(\rho)$. Then, the II estimator is defined as

$$\hat{\rho}_{II} = argmin_{\rho \in \Lambda} \mid \hat{\rho}_{OLS} - \frac{1}{K} \sum_{k=1}^{K} \hat{\rho}_k(\rho) \mid, \tag{9}$$

where $\hat{\rho}_{OLS}$ is the OLS estimator of ρ using the true dataset. Kyriacou et al. (2017) derive the asymptotic normality of the estimator without relying on a parametric assumption of the distribution of the disturbances, but requiring that they are iid and that there are no covariates other than the spatial lag of the dependent variable.

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