Local price spillovers in the Swedish District Heating sector

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Abstract

This study examines whether threat of stricter regulation influences the pricing behavior of firms. Using data on unregulated Swedish local district heating monopolists, we measure the threat level by the number of customer complaints about prices received by a national board. Exploiting a natural experiment, we find that firms homogenize their prices to reduce threat of regulation.

Keywords: regulatory threat, monopoly, price setting, spatial interaction, natural experiment.

JEL Codes: L12; L43.

1 Extended Abstract

We study the pricing behavior of Swedish District Heating (DH) firms. The Swedish DH sector represents a unique setup from an economic perspective. First, it consists of many local markets and in each market there is a single producer of DH. In addition, the vast majority of firms operates in only one market each. Second, the price of DH is not regulated. As a consequence, the researcher observes the unregulated pricing behavior of a large number of monopolists, each producing the same homogeneous good.

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We document sizable spatial correlation in DH prices. The first contribution of our paper is to study whether the spatial price correlation goes beyond co-movement of factors of supply and demand. In particular, we are interested in identifying price spillover effects. This is a difficult task because of endogeneity that arises from unobserved confounders and from the reverse causality relationship between prices. We address this problem with three different approaches that utilize different sources of identification. In our first approach, we use an innovative 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 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.

These spatial price spillovers pose a puzzle for several reasons. First, the monopolists in the DH sector are (largely) independent entities, i.e. they do not share profits, management, infrastructure or customers. Second, we put to test several mechanisms that appear to be natural candidates for explanations. In particular, firms might try to avoid large price differences to avoid (i) antitrust scrutiny and (ii) bad reputation among existing and potential customers. Our empirical tests reject both (i) and (ii).

To explain the spatial price spillovers, we suggest a novel mechanism that relates price differences to threat of stricter price regulation. Specifically, the DH price is not regulated but introducing regulation has been a subject of continuous political debate since 2005. Since 2008, the peak of these discussions, customers can complain about DH prices to a national DH board. This board has no legal power to impose restrictions on the prices of DH firms, but it makes their prices and customer complaints visible to the public. Bonev et al. (2019) provide ample anecdotal evidence that customer complaints to the national DH board are associated with increased likelihood of stricter regulation. In addition, the empirical findings of Bonev et al. (2019) suggest that complaints lead firms to decrease their DH prices, which is in line with the so-called Regulatory Threat Hypothesis (RTH), Glazer and McMillan (1992).

The mechanism suggested in this paper builds on the conjecture that customer complaints increase the likelihood of regulation. Our hypothesis is that firms homogenize prices across neighboring markets as a strategic move to avoid complaints. The underlying theory is that this strategy helps to convince customers that the local price is "fair", thereby reducing complaints. This theory draws on theoretical and empirical findings in behavioral economics, according to which customers might use prices in similar markets as reference prices, Kahneman et al. (1986), Rotemberg (2005) and Rotemberg (2011)). At the heart of the mechanism is the testable hypothesis, that when the price in the own market is much larger compared to reference prices, the customer will suffer an emotional cost and complain. We provide empirical support for this hypothesis. Thus, search costs of information, together with consumer utility generated from being treated fairly, generate an interaction of otherwise unrelated monopolists.

In a minor deviation of standard scientific practice, we then take a critical stand and thoroughly discuss the drawbacks of our analysis. One theoretical pitfall is the relationship of our mechanism to the Regulatory Threat Hypothesis. In particular, the latter appears difficult to derive in a context of industry-wide threat of regulation. Another pitfall is that in our simple model, an infinite number of equilibria are possible. Thus, it is hard to study welfare effects of threat. Our analysis is therefore to be understood as a first, modest, step towards a mechanism that matches empirical evidence, rather than a full-blown microeconomic analysis.

Our paper contributes to the literature on the effect of regulatory threat on prices, see section 4.1 for references. We complement this literature by (i) suggesting a behavioral mechanism that creates threat and (ii) an additional channel through which threat impacts prices.

The paper is structured as follows. In the next section, we briefly discuss the institutional setup and present the data and descriptive statistics. In section 3, identification and estimation of spatial price spillovers is discussed. Section 4 contains discussion about mechanisms, and section 5 about pitfalls of our analysis.

2 Institutional setup and data

2.1 Institutional setup

The exposition of the institutional setup is borrowed from our companion paper Bonev et al. (2019). Currently, there is one DH firm in the main locality in 262 of the 290 municipalities. 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.[1] 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.[2] For customers connected to the DH network, DH is the cheapest source of heat compared to electricity with a ratio of

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

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

variable costs around 0.5, EMI (2012). Due to the geographical restriction of DH and the high switching cost[3], 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 (transmission and local distribution) are regulated by the Swedish Energy Markets Inspectorate through ex ante revenue caps.

At the end of 2005, a national debate started with calls for regulation of DH prices (SOU 2005a, b). The regulatory debate culminated in the adoption of the District Heating Act (2008:263) in 2008. The most important provision in this law was the establishment of the Swedish District Heating Board. Since July 2008, a consumer who is dissatisfied about his/her DH price can file a complaint to the DH board. Notably, a customer can only complain to the DH board because of dissatisfaction with the price. Furthermore, dealing with the complaints about prices is the only function of the DH board. If the complaint is considered well-grounded, the committee launches a negotiation process with the consumer and the utility, and it provides expert opinions about how the DH price should be determined. The DH firm can accept or reject the board's suggestion without any direct consequences. Therefore, the District Heating Act provides consumers with no real additional rights, but it exposes consumer dissatisfaction to the public. Bonev et al. (2019) provide abundant anecdotal evidence that complaints to the DH board can be linked to threat of regulation.

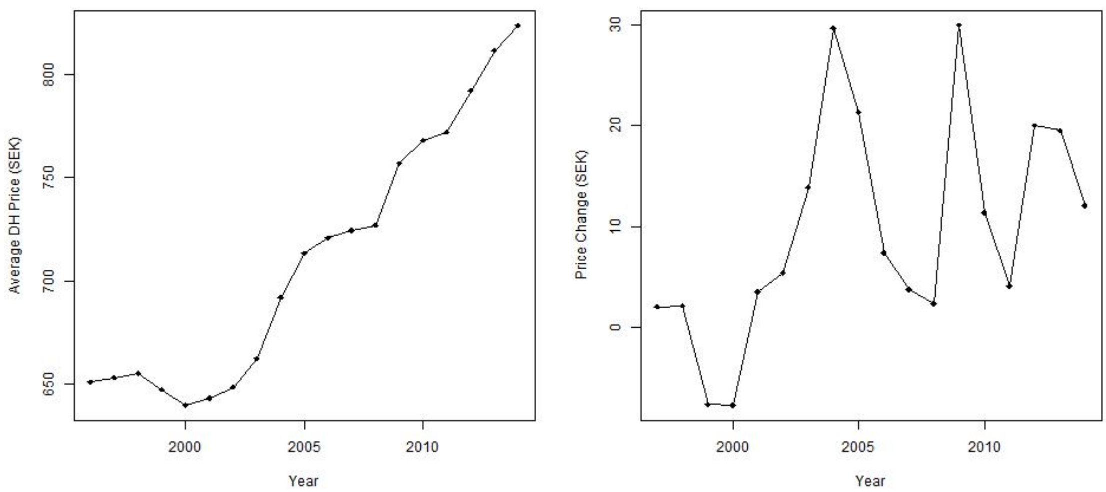
[3] A one-time connection fee is paid at the time when the residence connects to the network. This fee is high and can amount to ten times the total annual consumption cost.

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.[4] 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. After 2000, prices steadily increased. By 2005, when the debate about regulation started (the beginning of the threat period), prices had increased by a total of 11%. By the peak of the debate at the end of 2008, prices were almost 20% higher than in 2000. It is interesting to note that after 2005, prices increased at a slower pace. After 2009, price

[4] 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.

changes follow no clear pattern.

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

Table 1: Summary statistics of the sample used to test the SIH (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

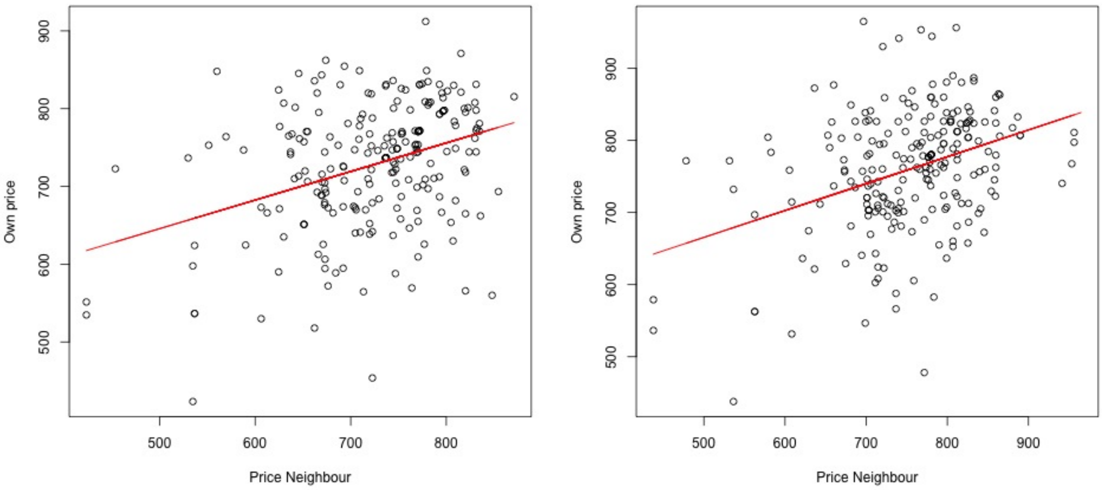
Note: Rows 3-8 contain statistics for the relative change (Rel. Δ) 2008-09 of a variable. Z_{i2009} is the local share of detached houses in 2009. Summary based on 225 observations.

we show descriptive statistics for the period 2008-2009. The choice for this time span is motivated in the identication 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 3.2. The relative price change, on the other hand, exhibits substantial variation, with largest price change being 28% and the lowest being negative. Figures 2a and 2b

Figure 2: Regression of DH in a market on the DH price in the closest neighbor market.

(a) Prices in 2008 (b) Prices in 2009



depict a regression of the price in a market on the price in the closest neighbor market in 2008 and 2009, respectively. These two prices appear to be highly positively correlated. This result remains valid if we take the correlation between the price in a market and the average prices of its 5 closest markets, or the average price of its direct neighbours. For now, we do not interpret this finding causally. Disentangling the different mechanisms behind it is the main objective of the paper.

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.[5] 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.

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 (a report on the Swedish media landscape in Sweden can be found on the webpage of the European Journalism Centre, http://ejc.net). Second, a big share of the working population commutes to one of the neighboring municipalities.[6] 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.[7]

3 Identification and estimation of price spillover effects

Define p_{it} to be the price charged by the local monopolist in market i at time t and let p_{-it} be the (weighted) average of prices in markets that are "close" to market i. A precise definition of and motivation for closeness will be given below. We are interested

[5] Descriptive statistics is available upon request.
[6] See e.g. http://www.grs.scb.se.
[7] This information is obtained from the Swedish Statistics Institute, http://www.scb.se/

in measuring the local price spillovers, i.e. the causel effect of p_{-it} on p_{it}. A major challenge for identification is that the spatial lag p_{-it} is potentially endogeneous. There are two main sources of endogeneity. First, prices are determined simultaneously and p_{-it} on p_{it} are thus potentially in a reverse causality relationship. Second, there could be unobserved confounders. In particular, the spatial price correlation documented in section 2.2 could be due to correlation of unobserved factors of demand and supply. 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. 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 50000 SEK (equivalent to about 4800 € 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 4 days later: 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 $t-1$.[8] 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.[9]. 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.

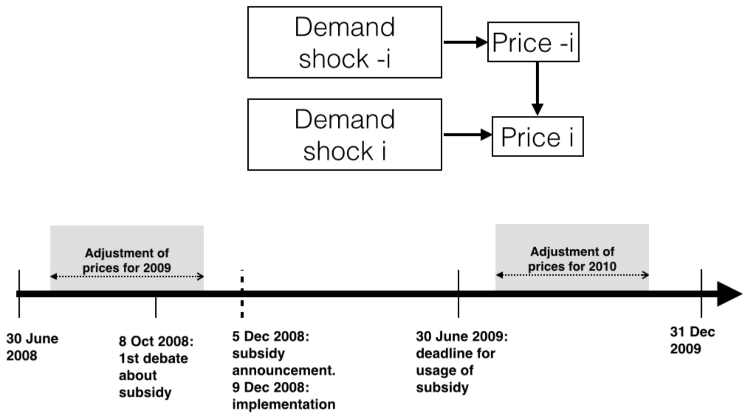
One effect of this subsidy was to make connecting to DH cheaper in 2009, which is likely to have attracted new consumers. Since the subsidy could only be taken until June 2009, this change in the stock of customers must have occurred prior to adjustment of prices for 2010. As a result, the subsidy is 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 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

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

[9] 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 .

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.



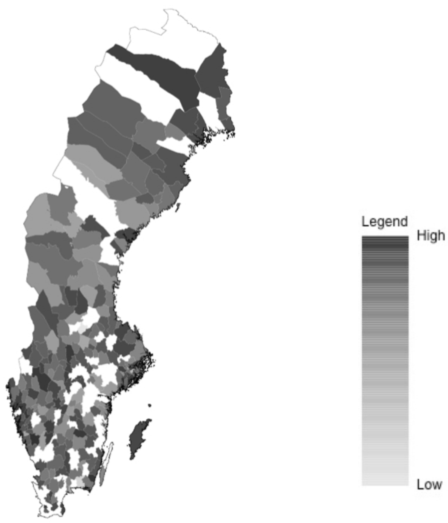
Unfortunately, we do not observe changes in the stock of customers triggered by the subsidy as we have no information on individual households' decisions to claim the subsidy. Instead, we use as a proxy the following variable

$$Z_i := \frac{\text{Number of households in detached houses in municipality i}}{\text{Total number of households in municipality i}}. \tag{1}$$

The interpretation of Z_{-it} is as follows. The nominator in (1) 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 (1) measures the total size of the demand of the DH firm in the fictive case where all households are customers. Thus, the variable Z_{it} 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} as an instrument for the endogenous price difference $p_{-i2009} - p_{-i2008}$, where Z_{-it} is defined analogously to the definition of p_{-it} and is equal to $\sum_{j \neq i} w_{ij} Z_{jt}$. In particular, 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.

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



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 on 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 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.[10] 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 is divided into 3 large "fuel regions". Within each of those regions, markets are exposed to the same price for most fuels, including biofuel (which is the major fuel type used by DH firms). Thus, variation in fuel cost is not related to local market characteristics.[11]

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 2. Although the estimated effect of

[10] 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.

[11] 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/

Table 2: 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 would assume that a similar correlation holds for unobserved factors. Thus, was the instrument related to these unobserved factors, one would expect to find spatial correlation in its distribution as well.

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}.

3.2 Estimation

We estimate the model

$$p_{it+1} = \beta_0 + \rho \sum_{j \neq i} w_{ij} p_{jt+1} + X_{it}\beta + \sum_{l=1}^{T-1} \delta_l T_l + \theta_i + \varepsilon_{it}, \tag{2}$$

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}, $\rho = \partial p_{it}/\partial p_{-it}$, with $p_{-it} = \sum_{j \neq i} w_{ij} p_{jt}$. It reflects the strength of the spatial spillover of neighboring prices on the price in market i. 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. This choice can be motivated with view of the possible explanations considered in the next section. In particular, we will consider mechanisms, in which customers compare prices in DH markets that are spatially close.

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

$$p_{i2010} - p_{i2009} = \beta_0 + \rho \sum_{j \neq i} w_{ij} (p_{j2010} - p_{j2009}) + (X_{i2009} - X_{i2008})\beta + (\varepsilon_{i2009} - \varepsilon_{i2008}). \tag{3}$$

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

Table 3 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 B 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 (3) 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),

$$\varepsilon_t = \eta M \varepsilon_t + \xi_t, \tag{4}$$

where M is a $n \times n$ spatial weights matrix and $\xi_t = (\xi_{1,t}, \dots, \xi_{n,t})$ is a vector of independent innovations with variances $\sigma_1, \dots, \sigma_n$. 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.[12] The IV estimates vary between 0.55 and 0.77, thus all lying within the 90 % confidence interval around 0.62 (the intermediate estimate). Since we instrument

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

Table 3: Empirical results, spatial price spillovers.

	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 $p < 0.01$.

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 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 A in the online appendix. 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.

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.

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.

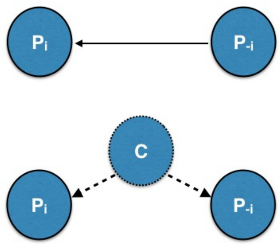
4 Explaining the spatial price spillovers: competing mechanisms

In this section, we put to test three competing explanations of the spillover effects. Our working definition of competing mechanisms is based on Card et al. (2011). Specifically, competing mechanisms yield a similar response with respect to manipulating the main independent variable, but a different prediction when another variable is manipulated.

The three mechanisms we study are regulatory threat, fear of antitrust scrutiny, and fear of losing potential customers due to bad reputation.

As a preliminary remark we note that the spatial correlation found in the previous section cannot be explained by spatial correlation of (observed or unobserved) supply and demand characteristics. In particular, the spatial spillovers are direct effects of prices on neighboring prices, and thus, due to the Reichenbach's principle of common cause, they cannot be due to merely common confounders (as depicted in Figure 5).

Figure 5: A direct effect vs. common confounders



4.1 Threat of regulation

Consider n markets in which a homogeneous good g (DH) is sold. In each market, a local monopolists serves consumers. Each consumer from market i either purchases one unit of g at the local price p_i, $i = 1, \ldots, n$, or switches to a substitute (electricity-based heating). Individual consumer utility for g in market i is u_i. The net utility of the outside option is normalized to zero. As a result, a consumer purchases g if $p_i \leq u_i$. The number of consumers in each market is normalized to $n_i = 1$.

Consumers in market i compare the price charged in that market, p_i, with prices in close markets. If p_i is too high compared to the average price in close markets, customers complain. This assumption draws on theoretical and empirical findings in behavioural economics. An important result in this field 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, in particular, the prices set by firms producing a similar type of goods or services, Kahneman et al. (1986), Rotemberg (2005) and Rotemberg (2011). If the price is judged as unfair, consumers suffer an emotional cost and might get angry. This adds a second, behavioural, component to the pure utility u_i of consuming the good. With these considerations, we assume that the number of complaints C_i in market i is generated according to

$$C_i = \psi(p_i, p_{-i}, A_i) \quad \text{with} \quad \partial\psi/\partial p_i > 0 \quad \text{and} \quad \partial\psi/\partial p_{-it} < 0, \tag{5}$$

A_i is a vector of factors influencing the propensity to complain and ψ is a function that relates the different variables. According to relationship (5), the higher the price in market i, the higher the level of customer dissatisfaction, which manifests itself in the number of complaints. In contrast, higher prices in surrounding markets reduce the number of complaints. The focus only on surrounding markets can be justified with information cost proportional to distance, see the remark "The local flow of information" at the end of section 2.2.

Next, let the local production cost be c_i, so that the service is provided only if $p_i \geq c_i$. We assume that $u_i \geq c_i$. The price is freely chosen by each firm. In addition, the monopolist in market i has knowledge of the customer utility u_i. Under *laissez-faire*, firm i would thus set the monopoly price $p_i^m = u_i$.

We further assume that there is a threat that a regulation law is enacted. The regulator is a perfectly informed body that maximizes consumer surplus. The regulated price would thus be equal to the marginal cost c_i. Following a long tradition in the theory of regulatory economics, we assume that enacting a new law is costly, see e.g. Stigler (1971). Legislation costs may be induced for example by the presence of lobby groups. We assume that the cost of regulation is a decreasing function of the complaints C_i, so that the firm has incentives to avoid complaints. The simplest way to model this relationship is to assume the following relationships:

$$p_i = \phi(C_i, V_i) \tag{6}$$

$$\frac{\partial \phi}{\partial C_i} < 0 \tag{7}$$

where V_{it} is a vector of other factors that influence the price and ϕ is a function that relates the different variables. If we interpret C_i as a proxy for the likelihood of introducing new regulation, then relationship (7) represents the Regulatory Threat Hypothesis. The RTH has been empirically studied by Erfle et al. (1989), Erfle and McMillan (1990), Wolfram (1999), Boyer (2000), Acutt et al. (2001), Stango (2003), Antweiler (2003), Ellison and Wolfram (2006) and Bonev et al. (2019). In the context of the Swedish DH sector, Bonev et al. (2019) provide (i) anecdotal evidence that complaints are associated with threat of regulation and (ii) empirical evidence in favor of (7). The mechanism, presented in this section, relies on both (i) and (ii), and we take them for granted. In section 5, we discuss the hidden assumptions behind (7) and try to relax it.

Plugging (5) in (6) yields

$$p_i = \phi(\psi(p_i, p_{-i}, X_i), A_i)$$

This equation defines an implicit relationship between p_i and p_{-i}. Assuming an exogenous shock on p_{-i}, we can differentiate this expression with respect to p_{-i}, leading to

$$\frac{\partial p_{it}}{\partial p_{-it}} \times \left[1 - \frac{\partial \phi}{\partial C_{it}} \frac{\partial \psi}{\partial p_{it}}\right] = \frac{\partial \phi}{\partial C_{it}} \frac{\partial \psi}{\partial p_{-it}}$$

Given the assumptions made above on the sign of different derivatives of ϕ and ψ, we immediately derive the following relationship:

$$\partial p_i / \partial p_{-i} > 0. \tag{8}$$

The intuition behind (8) is that an exogenous increase of p_{-i} decreases the complaints C_i, which gives a possibility for firm i to increase profits without triggering regulation. Thus, the threat of regulation generates interaction of otherwise unrelated firms. The interactions are "local" because of the information cost associated with price comparisons.

Remark Relationship 5 is testable. One difficulty associated with its empirical testing is that complaints and prices are in a reverse causality relationship. To account for the endogeneity of prices, we use an instrument on the supply side. In particular, we instrument prices with the firm-specific fuel mix used to generate heat. Since there is no cross-sectional variation in fuel costs, any change in the variable costs due to a shock in fuel prices is triggered only by the firm-specific ratio of using different fuel types. This cost shifter is used in the empirical IO literature as an instrument, see MacKay and Miller (2019) for a discussion and references. Given the large amount of findings presented in this paper, these results are left for the appendix. In summary, our results support the validity of 5, however they have to be interpreted with caution because of a weak instrument problem.

4.2 Threat of antitrust prosecution

Firms might reduce their prices when they fear antitrust scrutiny. In particular, the antitrust authority might interpret large differences in prices as a signal for violation of market power. XXX REFERENCE? Thus, fear of antitrust scrutiny could provide an incentive for firms to homogenize their prices.

Threat of antitrust action can be considered stable from 1996 until today. Moreover,

Table 4: Testing for spatial price spillovers under a No-Threat scenario.

	(1)	(2)
ρ	0.071	0.059
	(0.60)	(0.61)
Population		0.02
		(0.04)
Age > 65		1.85
		(1.70)
Labour Cost		
		(0.004)
Intercept	6.25	-1.68
	(12.81)	(13.57)
Num of obs.	215	215
I-Stage F-stat.	9.039	7.187

Note: price change is regressed on average DH price in neighboring markets. Data from 2004-2005. Equal weights for neighbors sharing a border, and 0 weight for all others. * denotes $p < 0.1$, ** denotes $p < 0.05$, *** denotes $p < 0.01$. One-sided test p-values for $\hat{\rho}$.

abuse of market power in has been a violation of the Swedish Competition Act (2008:579) since 1993. This institutional characteristic provides a source of discriminatory variation: if the spacial price spillovers are due to fear of antitrust prosecution, then we should be able to detect such spillovers at different points in time, most notably before the period of threat of regulation (i.e. prior 2005).

We estimate the spatial spillover effect ρ from (2) using data from the period 2004-2005. In 2004, a policy shock analogous to the 2008-subsidy took place.[13] We use this policy to identify the price spillover ρ. Table 4 presents the results obtained under a SARAR(1,1) model with weights equal among all neighbors sharing a border with market i (and otherwise 0). In the first regression, only the spacial price lag is included as a covariate, whereas regression 2 includes two additional covariates (we do not observe labor cost for this period). The estimates of ρ are close to 0 and insignificant in both specifications.[14] These insignificant results are not compatible with time-steady fear of

[13] A description of the subsidy is provided in section D in the online appendix.
[14] Alternative specifications yield similar results and are omitted.

antitrust intervention being the driving force behind the spatial price spillovers.

4.3 Threat of losing potential customers

Positive local correlation of prices could be generated by the firms' objective of attracting new customers (existing customers are locked in). Potential customers who perceive the DH price as unfair might opt for alternative heating options (Rotemberg (2011)). This possibility may lead firms to reduce price differences across districts, just as they would in the case of threat of regulation. In the context of threat of regulation, this alternative mechanism was first discussed by Olmstead and Rhode (1985). In order to disentangle the effect of a regulatory threat from the threat of losing potential customers, we construct a measure of potential loss of customers. We take the (relative) increase in the number of constructed dwellings in a municipality during the coming three years, i.e. 2010-2013. The future growth in dwelling stock of a municipality is strongly correlated with the number of potential customers: in principle, each new dwelling is occupied by a new potential customer. Moreover, the short and middle term growth in the number of dwellings in a municipality is known in advance, as building permissions are typically granted 2-3 years before the completion of a building. We observe this growth for the years 2010-2013 and assume that it was perfectly known to the firms in 2009.[15] We then estimate ρ using the 50% of municipalities with the highest growth in dwelling stock. If the competition mechanism is responsible for the spatial price spillovers, we should see a (substantially) higher estimate of ρ for this subsample than for the full sample.

The results of two different specifications are shown in Table 5. The first specification contains only the neighboring price as a covariate, and the second specification includes all observed covariates. The estimates $\hat{\rho}$ from the two specifications are very close in magnitude to the corresponding estimates when we use the full sample (i.e. in an estimation with all markets, as in table 3), and are also significant. Thus, these results

[15] Building permissions can be observed by any citizen or organization and DH firms obviously have strong incentives to stay updated about construction plans of dwellings.

Table 5: Testing for spatial price spillovers with the subsample of the 50 % municipalities with highest future growth.

	(1)	(2)
ρ	0.60**	0.51*
	(0.35)	(0.37)
Population		0.0003
		(0.0002)
Age > 65		-1.36
		(0.88)
Labour Cost		-0.002
		(0.008)
Intercept	11.81	7.87
	(11.72)	(16.03)
Num of obs.	114	114
I-Stage F-stat.	9.907	11.41

Note: main results SIH. Price change is regressed on average price of neighbors. Specification (1) assumes independent disturbances, the standard errors are heteroskedasticity robust. Specification (2) is a SARAR(1,1) model. Specification (3) includes observed covariates. Weights are 1 for neighbors with shared border and zero otherwise (row standardized). * denotes $p < 0.1$, ** denotes $p < 0.05$, *** denotes $p < 0.01$. One-sided test p-values for $\hat{\rho}$.

are not consistent with the competition mechanism.

5 Pitfalls

In this section, we discuss potential pitfalls of the threat mechanism. The first problem revolves around the local vs. global nature of threat. The second one is that there are infinitly many equilibria. The third one concerns the indirect nature of our results. To fix ideas, denote $R \in \{0, 1\}$ to be the decision of the regulator to regulate ($R = 1$) or not ($R = 0$). Furthermore, denote by K the cost of regulation. To simplify the discussion, we assume that there are only two firms, $n = 2$.

Local vs. Global Regulation. In section 4.1, we directly assumed $\partial p_i/\partial c_i < 0$ and interpreted C_i as the *firm-specific* threat of regulation. The validity of this interpretation, however, depends on the assumptions about the nature of threat. In particular, consider the following two cases.

Case 1 (local threat): the regulator decides case by case to regulate or not. In each market-specific decision, only the market-specific costs and benefits are considered. Formally, the decision of the regulator can be written as

$$R_i = 1 \quad \text{iff} \quad K_i(C_i) \leq p_i - c_i. \tag{9}$$

The r.h.s. of the inequality above represents the gains from regulating market i. In this setup, $\partial p_i / \partial C_i < 0$ is straightforward to derive.

Case 2 (global threat): the decision of the regulator is to regulate or not the whole industry. The cost of regulation depends on all complaints, $K = K(\sum_{i=1}^{n} C_i)$. Here, an increase in the complaints in market i has an impact on all firms in the industry and it is not clear what the individual best response should be. We note, however, that under the assumption that the political costs react sensitively to a change in complaints, one can still derive $\rho > 0$. To state this formally, we need the following definitions. Set $c = (c_1 + c_2)$. Furthermore, define $\Phi = K \circ F$ and let $b_i(p_{-i})$ be the best response of player i to a price p_{-i} played by the other player. Define the function G as

$$G(x, y) = \Phi(\mid y - x \mid) - x - y + c. \tag{10}$$

Then, ignoring the index $-i$, the best response function b is an implicit function of p defined by

$$G(p, b(p)) = 0. \tag{11}$$

In the following result, we implicitly assume that participation constraints of the firms are satisfied and that the presence of threat is indeed binding for the firms (i.e. K are not too high).

Proposition 5.1. *Suppose that $F' > 0$ and $K' < 0$ and suppose that for some interior p the indifference condition* 11 *is fulfilled. Denote $b(p) = \eta$. Then* 11 *holds for a* *neighbourhood around p. Moreover, if in addition*

$$|\Phi'| > 1, \tag{12}$$

then the derivative of b satisfies $b' > 0$.

The proof is in the appendix. Since per definition $\Phi = K \circ F$, then (12) is satisfied when K' is very large. A large K' reflects a sensible environment, in which a small number of complaints are associated with a large change in the political incentives to regulate. Given the heated political und public debate in the period which we use for estimation, assumption (12) appears to be plausible.

As a result, our simple model predicts a positive spatial price spillover coefficient $\rho = b'$ for both a setup with local threat and a setup with global threat in a sensitive environment.

Equilibrium behavior. Assume now we are in a setup with global threat (case 2). Assume further that for $i = 1, 2$, $c_i \leq p_i \leq u_i$ (participation constraints). It is trivial to show, that any pair (p_1, p_2) that satisfies

$$p_1 + p_2 - c = K(F(p_1 - p_2) + F(p_2 - p_1)) \tag{13}$$

is an equilibrium. Thus, under mild assumptions on F and K, there are infinitly many equilibria. As a result, our framework has a limited value of studying the effect of threat on firms *in equilibrium.*

Indirect nature of threat. A final problem in our study is the indirect nature of the proxy for threat. In particular, since $R = 1$ is never observed, the claim that customer complaints are a proxy for threat remains an argument for which we can provide anecdotal evidence at best. Precluding several alternative interpretations of C_i, as we did, is of limited value, see Card et al. (2011) for a discussion. This drawback,

Table 6: First stage results corresponding to the IV estimation in Table 3, section 3.2 in the main paper.

	(1)	(2)	(3)	(4)
Z_{it}	33.12*** (7.395)	35.30*** (8.16)	30.46*** (12.45)	33.12*** (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 3, section 3.2 in the main paper, respectively. * denotes $p < 0.1$, ** denotes $p < 0.05$, *** denotes $p < 0.01$.

however, is shared (even in a greater extend) by all empirical studies on regulary threat. One potential way to reduce the ambiguity of customer complaints as a threat generating mechanism is to elicit managers' and politicians' subjective probabilities of regulation via e.g. a survey methodology. This remains within the scope of future research.

A Estimating spatial price spillovers

In table 6, we present the first stage results corresponding to estimations (1)-(4) in table 3 in the main text.

B 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 $y_1(\rho), \ldots, y_K(\rho)$, 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

$$\hat{\rho}_{II} = argmin_{\rho \in \Lambda} \mid \hat{\rho}_{OLS} - \frac{1}{K} \sum_{k=1}^{K} \hat{\rho}_k(\rho) \mid, \tag{14}$$

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.

C Testing the complaints generating mechanism

To test the assumption (5) from the main paper, we specify the model

$$C_{it} = \alpha_0 + \alpha d_{it-1} + X_{it}\delta + \sum_{l=1}^{T-1} \tau_l T_l + \eta_i + \nu_{it}, \tag{15}$$

where d_{it} is defined as $p_{it} - \sum_{j \neq i} w_{ij} p_{jt}$. It gives the difference between the price set by firm i and the weighted average of the prices of its neighbours. Thus, we expect α to have a positive coefficient, and define the Null hypothesis $H_{0,A} : \alpha \leq 0$.

To account for potential endogeneity of d_{it}, we use two cost shifters as instruments. Since C_{it} is determined on the demand side, a cost shifter would be a valid exclusion restriction. This line of reasoning follows the standard IV identification in the demand estimation literature. We use (i) the share of wasted heat of the total annual amount of fuel used by a DH firm and (ii) the share of oil on the total annual amount of fuel as instruments. Waste heat is a byproduct in industrial processes and sometimes it can be fed into the DH system to provide a cheap source of complimentary heat. Oil, on the other hand, was the most common fuel type in Swedish DH systems until the 1980s. Now the average share is less than 15% and it is primarily used as startup fuel and to meet peak demand during cold periods. The marginal cost of oil is relatively high and

DH firms are likely to adjust their prices whenever using higher shares of it.

Table 7 presents summary statistics of all variables used in our regressions. Since we do not use the demand shock as an instrument here, we are not restricted to using only observations from 2008-2009. Thus, our sample contains observations for the whole period 2009-2014 for which complaints data is available.

Table 7: Summary statistics of of the sample used to test equation 5 (2009-2014)

Variable	Min	1st Qu	Median	Mean	3rd Qu	Max
Price	536.5	743.61	795.9	786.17	839.63	1163.23
Population	1017	4794	8464	77 927	21 494	1469131
Labor cost	23 089	24 631	25 200	25 363	26 000	30 744
Age > 65	0.13	0.19	0.22	0.22	0.24	0.32
Complaints	0	0	0	0.0731	0	81
Share fuel 1	0	0	0	0.071	0.033	0.99
Share fuel 2	0	0.0001	0.0007	0.0022	0.0023	0.15

Note: fuel 1 = fuel from waste heat, fuel 2 = oil fuel. Summary based on 1519 observations.

Table 8 presents the results of four different specifications of (15). Specifications (1) and (2) are obtained under equal weights for all bordering neighbors, whereas (3) and (4) under weights proportional to the distance to the bordering neighbors. The instruments used to produce the results in Table 8 are Share of heat, Share of Oil and Share of Oil squared. We use the limited maximum likelihood (LIML) estimation method due to its better properties with weak instruments.

The estimates of α are very similar in all four regressions. They are positive and significant at the 10% level (p-values are calculated for a one-sided test corresponding to the hypothesis tested). An increase of 75 SEK (roughly 10% of the average price) would increase the number of complaints in a market by 0.14 complaints. This appears at first to be a very small number. We note however that the average number of complaints per

Table 8: Empirical results, testing the complaints generating hypothesis (5).

	(1)	(2)	(3)	(4))
Price difference	0.0014*	0.0014*	0.0015*	0.0015*
	(0.0011)	(0.0011)	(0.0011)	(0.0011)
Population		-1.08e-06*		-1.09e-06*
		(5.87e-07)		(5.82e-07)
Age > 65		0.763		0.796
		(0.939)		(0.955)
Labour Cost		0.000		0.000
		(0.00001)		(0.000014)
Year dummies	Yes	Yes	Yes	Yes
Fixed effects	Yes	Yes	Yes	Yes
Num of obs.	1519	1519	1519	1519
I-Stage F-stat.	5.23	5.15	4.76	4.71

Note: Specifications (1) and (2) are obtained under equal weights for all bordering neighbors, whereas (3) and (4) under weights proportional to the distance to the bordering neighbors. HAC-robust errors in all specifications. * denotes $p < 0.1$, ** denotes $p < 0.05$, *** denotes $p < 0.01$. One-sided test p-values for the Price difference variable.

municipality and per year is 0.19. Therefore, an increase of 0.14 complaints corresponds to more than 10 % of the average number of complaints per municipality for the whole period, and more than 50% of the average number of complaints per municipality and per year. This finding provides evidence that customers use neighboring prices as reference prices in order to learn about the fair price of DH. Big positive difference to neighboring prices is considered unfair and penalized with complaints.

D Description of the energy efficiency subsidy in 2004-2005

The Swedish Government decided on the 15 April 2004 to offer a subsidy that homeowners could use for all repair, conversion and extension works that improved the energy efficiency of their houses. The subsidy could be claimed from the 15 April 2004 until the 31 June 2005. No Parliamentary debate proceeded the implementation of the subsidy, but the Government presented and implemented the subsidy on the same day, i.e. 15 April. The maximum amount home-owners could receive was 30% of the labor cost, or 10500 SEK for single family (detached) dwellings and 5000 SEK for apartments in attached buildings.

E Proofs of theoretical results

Proof 1 (Proof of Lemma 5.1)**.** *Using the Implicit Function Theorem, we have*

$$\begin{aligned} b'(p) &= -\frac{\partial G(p,b(p))/\partial x}{\partial G(p,b(p))/\partial y} \\ &= -\frac{K'(F(p-b(p))+F(b(p)-p))(F'(p-b(p))-F'(b(p)-p))-1}{K'(F(p-b(p))+F(b(p)-p))(F'(b(p)-p)-F'(p-b(p)))-1} \\ &= 1+\frac{2}{K'(F(p-b(p))+F(b(p)-p))(F'(b(p)-p)-F'(p-b(p)))-1}. \end{aligned}$$

In order to assess the sign of b', we differentiate between the following cases.

Case 1: $p > b(p)$. Then $F(b(p)-p)=0$. Consequently, $b'(p) > 0$ is equivalent to

$$K'(F(p-b(p)))F'(p-b(p)) > 1. \tag{16}$$

Case 2: $p < b(p)$. Then $F(p-b(p))=0$, and $b'(p) > 0$ is equivalent to

$$K'(F(b(p)-p))F'(b(p)-p) < -1. \tag{17}$$

Therefore, in both cases $b'(p) > 0$ is equivalent to $\mid \Phi(\mid p-b(p) \mid) \mid > 1$.

References

Acutt, M., Elliott, C., and Robinson, T. (2001). Credible regulatory threats. *Energy Policy*, 29(11):911–916.

Anselin, L. and Florax, R. (1995). *New Directions in Spatial Econometrics.* Springer Berlin Heidelberg.

Antweiler, W. (2003). How effective is green regulatory threat? *American Economic Review*, 93(2):436–441.

Biggar, D., Glachant, M., and Söderberg, M. (2018). Monopoly regulation when customers need to make sunk investments: evidence from the Swedish district heating sector. *Journal of Regulatory Economics*, 54(1):14–40.

Bonev, P., Glachant, M., and Söderberg, M. (2019). Testing the regulatory threat hypothesis: Evidence from sweden. Working paper, University of St. Gallen.

Botterud, A., T., K., and Ilic, M. D. (2010). The relationship between spot and futures prices in the Nord Pool electricity market. *Energy Economics*, 32(5):967 – 978.

Boyer, M. M. (2000). Media attention, insurance regulation, and liability insurance pricing. *Journal of Risk and Insurance*, 67(1):37–72.

Brännlund, R., Ghalwash, T., and Nordström, J. (2007). Increased energy efficiency and the rebound effect: Effects on consumption and emissions. *Energy Economics*, 29(1):1 – 17.

Card, D., DellaVigna, S., and Malmendier, U. (2011). The role of theory in field experiments. *Journal of Economic Perspectives*, 25(3):39–62.

Difs, K. and Trygg, L. (2009). Pricing district heating by marginal cost. *Energy Policy*, 37(2):606 – 616.

Ellison, S. F. and Wolfram, C. (2006). Coordinating on lower prices: pharmaceutical pricing under political pressure. *RAND Journal of Economics*, 37(2):324–340.

EMI (2012). Uppvarmning i Sverige. Report EI 2012:09, Swedish Energy Markets Inspectorate.

Erfle, S. and McMillan, H. (1990). Media, political pressure, and the firm: The case of petroleum pricing in the late 1970s. *The Quarterly Journal of Economics*, 105(1):pp. 115–134.

Erfle, S., McMillan, H., and Grofman, B. (1989). Testing the Regulatory Threat Hypothesis with media coverage of the energy crisis and petroleum pricing in the late 1970s. *American Politics Quarterly*, 17(2):132–152.

Fuchs-Schuendeln, N. and Hassan, T. A. (2015). Natural experiments in macroeconomics. Working Paper 21228, National Bureau of Economic Research.

Glazer, A. and McMillan, H. (1992). Pricing by the firm under regulatory threat. *The Quarterly Journal of Economics*, 107(3):pp. 1089–1099.

Heckman, J. J., Lalonde, R. J., and Smith, J. A. (1999). The economics and econometrics of active labor market programs. In Ashenfelter, O. and Card, D., editors, *Handbook of Labor Economics*, volume 3, chapter 31, pages 1865–2097. Elsevier.

Kahneman, D., Knetsch, J. L., and Thaler, R. (1986). Fairness as a constraint on profit seeking: entitlements in the market. *American Economic Review*, 76(4):728–41.

Kelejian, H. H. and Prucha, I. R. (1998). A generalized spatial two-stage least squares procedure for estimating a spatial autoregressive model with autoregressive disturbances. *The Journal of Real Estate Finance and Economics*, 17(1):99–121.

Kyriacou, M., Phillips, P. C. B., and Rossi, F. (2017). Indirect inference in spatial autoregression. *The Econometrics Journal*, 20(2):168–189.

Lyytikäinen, T. (2012). Tax competition among local governments: Evidence from a property tax reform in Finland. *Journal of Public Economics*, 96(7-8):584–595.

MacKay, A. and Miller, N. (2019). Estimating models of supply and demand: Instruments and covariance restrictions. *Available at SSRN 3025845*.

Olmstead, A. L. and Rhode, P. (1985). Rationing without government: The West Coast gas famine of 1920. *The American Economic Review*, 75(5):1044–1055.

Rotemberg, J. J. (2005). Customer anger at price increases, changes in the frequency of price adjustment and monetary policy. *Journal of Monetary Economics*, 52(4):829–852.

Rotemberg, J. J. (2011). Fair pricing. *Journal of the European Economic Association*, 9(5):952–981.

Sjödin, J. and Henning, D. (2004). Calculating the marginal costs of a district-heating utility. *Applied Energy*, 78(1):1 – 18.

Stango, V. (2003). Strategic responses to regulatory threat in the credit card market. *The Journal of Law and Economics*, 46(2):427–452.

Stigler, G. J. (1971). The theory of economic regulation. *The Bell Journal of Economics and Management Science*, 2(1):3–21.

Wolfram, C. D. (1999). Measuring duopoly power in the British electricity spot market. *American Economic Review*, 89(4):805–826.