Competition and Product Price Dynamics in the Finnish District Heat Markets

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Abstract

Geographically household heat product markets are limited to the area which is covered by the local district heating network. A relevant case exists where the district heating company can determine the price of its product without constraints from other firms and heat products. We test empirically whether the local prices of district heating are affected by the local heat product market shares of district heating companies. In addition we test the price effects of public ownership of plants and production shares of district heat companies. We use panel data which consists of 76 district heating companies in years 1996 – 2002. The data includes market shares, joint production, district heating tariffs, production scale, and raw material input cost variables.

The results indicate that pure competitive case is not prevailing in the Finnish district heating pricing. The production shares of district heat firms had generally a positive effect on the district heating price but shares in the local heat product markets did not affected price. Public ownership was not relevant factor in pricing. Estimates from proposed model of price dynamics showed that some competitive patterns are quite strong among the market participants. Firms' pricing movements were sensitive - albeit stability preserving - to changes in other firm's pricing decisions in preceding periods. Generally the results indicate that the Finnish district heat markets are non-competitive but entail price dynamics that are typical for stable oligopolistic markets.

Keywords: Household energy markets, price setting, market shares

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I. INTRODUCTION

The Finnish heat product markets consist mainly of electricity, district heating, light and heavy fuel oil, and fire wood. The market shares of these products were following in 2003: district heat 49 %, light fuel oil 17 %, electricity 17 %, fire wood 12 %, heavy fuel oil 1 % and others 4 %.[1] The market definition of a different heat product differs. For example electricity, light and heavy fuel oils each form a single market in Finland. But, the market of a district heating company is limited to the area which its district heating network covers. From the district heating company's point of view heat product markets are limited to the area which is covered by the local district heating network and its competitors are other heat product offered in this area.

Do also the household customers feel this way? The electricity markets have been deregulated since November 1998. Still, the local electricity network operator has the ultimate service obligation to the customers located its network area. The amount of household electricity users who have changed their electricity supplier is quite low only 5 %. About 22 % have asked offers from other supplier but 17 % is still relied to the original supplier.[2] Therefore, the electricity markets are not quite competitive after all. The amount of competition in light fuel oil markets depends on the amount of suppliers in the area, but compared to other heat products these markets are most competitive.

Each of the heat products also demand different technology to be utilised. After the choice of a technology have been done it is not economically efficient to change between the heat products until the technology used wears out. Because, these heat product markets exist only in situations where the investment decision of a heating technology is done, the existence of these markets is difficult to catch trough empirical analysis. At the moment, the markets of household heat products are defined as independent of each other. In the article we test first whether district heat firm's market share in the *total heat product* markets has effects on its pricing. This enables us to test indirectly whether this market independency is valid, i.e. do different household heat products act as substitutes

[1] Energiateollisuus ry, (2005).
[2] Lindberg, Koskenrouta, & Timonen, (2003),16.

to each other. If substitution takes place then market shares should correlate with product prices. However, the product substitution may quite often be limited since the local district heat supplier is the only supplier on the area and also electricity companies have high market shares in the area they are located. The amount of competitors even in these enlarged markets is low.

This observation raises the question whether the local district heating network gives a technological potential to non-competitive product specific pricing. Thus, a relevant case exists where the district heating company can determine the price of its product without constraints from other firms and heat products. Non-competitive local markets may be dominated by few or even by a single firm. Market share effect of *district heat product* markets on pricing is tested also in addition to plant scale and unit cost effects.

Another interesting point is the amount of public ownership within the district heating companies. Since, the privatisation has entered to the energy markets also great amount district heating companies has been privatised. About half of the companies in the sample are in public ownership. Previous studies have illustrated that state owned enterprises are often instructed to pursue goals other than profit maximization such as setting prices below cost in order to increase local employment or to ensure affordable services to all customers. [3] Some of the management of a state owned utility might also be encouraged to maximize the amount of sales and size of a company in order to attain more appreciation. Therefore, we also tested whether the public owned companies pricing differ from the pricing of other companies.

In the empirical analysis we use panel data, which consists of 76 district heating companies in years 1996 – 2002. The data includes variables like the market share, the production technology (separate or CHP), the district heat prices, the scale of production, and raw material input costs.

The pricing of district heat differs strongly between companies with different market shares within the local heat product markets. The following two figures will give a slight

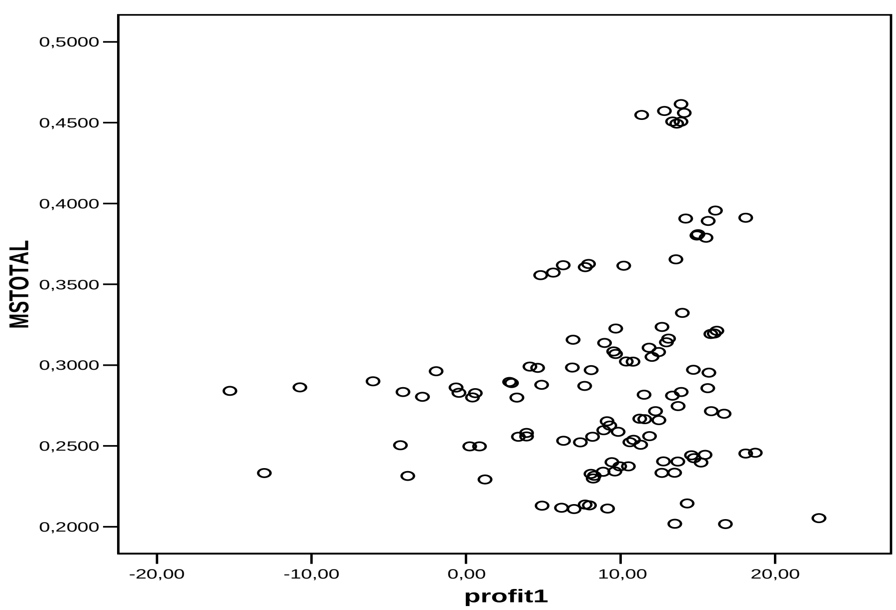
[3] Sappington & Sidak, (2003a), 479 – 480.

insight to the data. The first figure illustrates the situation in companies with 20 % or higher market share in the local heat product markets.

Table 1. Summary statistics of data: means and standard deviations (in parenthesis**)**

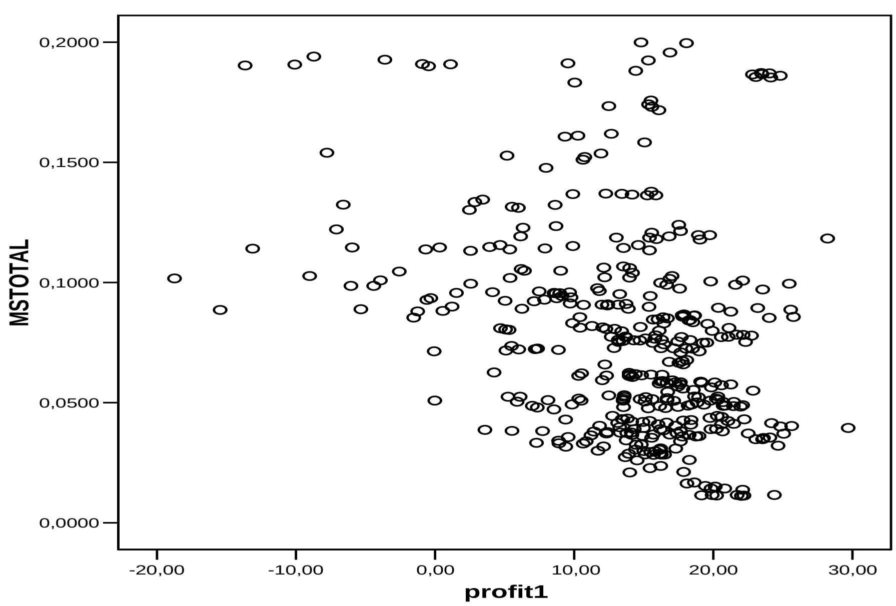
	1996	1997	1998	1999	2000	2001	2002
Average market share of district heat producers	63 % (42.9 %)	62 % (43.8 %)	63 % (42.2 %)	63 % (41.5 %)	62 % (42.4 %)	64 % (42.1%)	65 % 42.6 %
Total market share in heat product markets	12.4 % (10.1 %)	12.4 % (10.2 %)	12.4 % (10.2 %)	12.4 % (10.1 %)	12.4 % (10.1 %)	12.5 % (10.1 %)	12.6 % (10.2 %)
Market share in blocks and terraced houses heat prod. mkts	51.3 % (18.4 %)	51.5 % (18.3 %)	51.7 % (18.3 %)	52.4 % (18.1 %)	52.5 % (18.2 %)	53 % (17.9 %)	53.3 % (18 %)
Number of public own companies	35	34	33	33	33	33	33
Energy tariff (€/MWh)	25.6 (4.21)	26.5 (4.26)	27 (4.2)	26.6 (3.91)	28 (4.81)	31.3 (6.89)	30.5 (5.43)
Total tariff for blocks and terraced houses	38.7 (4.99)	39.5 (4.94)	40.2 (5.12)	39.5 (4.71)	41 (5.22)	44.2 (7.22)	44 (5.36)

Figure 1. The difference between energy tariff and estimated fuel costs in district heating companies which have 20 % or higher market share in the local heat product markets.



The figure gives a quite strongly increasing relation with profits and market share in the companies which market share is about 30 % or more. An interesting group is companies between 15 and 25 % market share. Their pricing or profit formation does not seem to have a straight pattern. It might be that these companies are large enough to attend the local heat product markets seriously and the profitability and pricing depends highly on the situation in the local markets. Some markets might be highly competitive and the price is set according to the marginal costs. Some publicly own companies might carry out pricing under marginal costs to maximize the use of district heating in the area. And in some markets the price level might be so high that district heating companies can earn high profits. Another interesting point is that district heating companies with less than 10 % market share in the local heat product markets also has possibilities to earn high profits and that when market share decreases the profits increase. It might be that these companies have group of loyal customers who they can rely on and then they can maximize profits through production technology or pricing since the demand is almost constant or the elasticity of demand is high.

Figure 2. The difference between energy tariff and estimated fuel costs in district heating companies with less than 20 % market share in the local heat product markets.



Finally our target is to analyse in details the product price dynamics between the local district heat companies. Although some firms may even be a single district heat producer in local markets their pricing policy is not independent of their local and non-local rivals. Mergers, entry and exit in local market have been quite active in Finnish heat markets since the mid 1990's. Strategic pricing decisions and dynamics of market structure depends on competitors' actions. Few producers on heat markets with similar plant cost structure refer to oligopolistic markets where product price correlation is high. Thus we test in details whether firms' pricing depends on their rivals' prices, price volatility, and changes in market shares.

II. MODELLING PRICE SETTING AND PRICE DYNAMICS

2.1. Background

The empirical literature of industrial economics faces many practical problems albeit the theory of competitive and non-competitive markets is well established. The theory of relevant market is a nice example of this disparity. Basically the notion that low number of producers is located in the same geographic market producing different or identical varieties of the same product is all that the theory gives (Stigler & Sherwin 1985). However, empirically the problem of discovering the relevant geographical market is at least as vexing as that of discovering the relevant product market. Price discrimination allows oligopolistic or monopolistic firms to invade their competitor's markets areas, so that the high degree of overlap and cross-hauling tends to obviate the usefulness of theory literature for practical applications (Elzinga & Hogarty 1973, Thisse & Vives 1988). In practice, the best guide is still good common sense and a detailed institutional knowledge (Kirman & Philps 1993).

2.2. SCP –paradigm

The case of district heat production is a quite easily defined as a relevant market with one commodity produced by low number of producers on separate geographic area. Basically we have to concentrate on the price data with Cournot's principle: market prices should take the same level throughout with ease and rapidity. The empirical implementation focuses on the dynamics of prices and exogeneity of price setting. Thus if price adjustment is rapid and other firms' prices have small effects on particular firms price setting the market regarded as competitive rather than oligopolistic (Slade 1986, Spiller & Huang 1986, Bresnahan 1989). The relevant literature focuses on the notion of conjectural variation where theory predicts that one unit change in firm's output leads to a change of γ_i in its competitors' total output. Formally we have when allowing for price-setting conduct that

$$p_t = c_{it} - \frac{dp_t}{dq_t} q_{it}\theta_{it} \tag{1}$$

where $p_t = p(\sum_i^N q_{it})$ is the market demand, c_{it} measures firm's marginal cost, and $\theta_{it} = 1 + \gamma_i$ gives the measure of the competitiveness of oligopoly conduct. The case $\theta_{it} = 1 \quad (\gamma_i = 0)$ is implied by Cournot-Nash equilibrium in a one-shot game (or monopoly). $\theta_{it} = 0 \quad (\gamma_i = -1)$ indicates perfectly competitive behaviour (Waterson 1984, Kirman & Philps 1993).

The structure-conduct-performance (SCP) paradigm has been a much used workhorse in the empirical industrial economics (see e.g. Reid 1987, Cubbin 1988, Waterson 1984). It has been under hard pressure for decades but still many found it as a valid starting point for empirical market share and pricing analysis (Barla 2000, Slade 2004). Basically the paradigm states that after rearranging Eq. 1) a following relationship for firm's price margin is relevant

$$PC_i = \frac{p - MC_i}{p} = \frac{S_i(1+\gamma_i)}{\eta} \ , \tag{2}$$

where p the market price of industry's output, MC_i is the ith firm's marginal costs, S_i is the firm's market share of output, and η is the industry price elasticity of demand.

Sappington and Sidak (2003) have constructed a theoretical model of public companies behaviour, assuming that company values both total incomes and profits. Public utility company maximizes following revenue function:

3) $$R(Q) = w\left[\sum_{i=1}^{n} p_i Q_i(p)\right] + [1-w]\left[\sum_{i=1}^{n} p_i Q_i(p) - C(Q)\right].$$

Where *w (0,1)* is the weight that public company gives to the income, *n* is the amount of goods supplied, $p_i \geq 0$ is company's price for product and $Q_i(p)$ is customers demand of product *i* at price *p*, which is set by company. The first term of the equation describes company's total income i.e. price multiplied with quantity and the second term company's profits. If *w = 0* the equation describes a situation in a normal private profit maximizing company.

When a normal profit maximising multi-product company sets price as was illustrated in Eq. 3.), a public multi product company sets price as follows:

4) $$PC_i = \frac{p_i - [1-w]\dfrac{\partial C_i(Q)}{\partial Q_i}}{p_i}, \text{with all } i = 1,..., n.$$ [4]

The pricing principle of public multi-product company is similar as private except for the fact that marginal costs are decreased with factor 1-*w*, which describes reduced aim to maximize profits. It follows that the public company favour pricing below costs, when it emphasizes more total sales and when demand is elastic. But then, if demand elasticity is constant, the public company sets the price below the marginal costs to all of the products with demand elasticity above 1/*w*. So, if *w* is 0.5 all products which have price elasticity 2 or more are priced below costs.[5]

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[4] Sappington & Sidak, (2003a), 504.
[5] Sappington & Sidak, (2003b), 192.

From the econometric point of view relations described in Eqs. 2) - 4) are problematic, since all variables are more or less endogenous. They all form the market process simultaneously making consistent empirical estimation difficult. In this context we will not model district heating companies' profits since we do not have direct observation for them. Instead we decompose Eq. 2) in following form

$$P_i = g(MARKET\ SHARES_i, C_i, X_i) \tag{5}$$

where P_i is the product price of district heating company, $MARKET\ SHARES_i$ contains both variables for firm's market share in *local heat markets* and output product share in *local district heat markets*, C_i is the (unit) cost for the firm, and X_i includes all other relevant variables that data allows for.

2.2.1. Econometric model

In order to analyze firm's market share effect on its price setting a following dynamic empirical panel data model is suggested (see Eq.6.). The model is a dynamic two way fixed effect model. The main interest lies in the effects of firm's market share of heat markets ($lnMS$) and in firm's production share of district heat *(lnPRODS)* on market tariff *(lnEprice)*. If the firm has a market power then the prediction is that shares correlate positively with price levels. Also we pay attention to public owned plants. $DPUB$ is a $0/1$ -dummy for the local public or state ownership of the firm. Firm's output price level is determined also by many other variables. The main cost component of heat energy production is the raw material costs. The variable $lnFUEL$ is a proxy for the not revealed firm's material input unit cost. The variable is constructed from market prices of raw materials (e.g. oil and gas) and total usage of firm's raw material inputs. The variable $lnSCALE$ describes the firm's output capacity, i.e. how much the firm has produced energy per year (measured in GKW). Variable $DJoint$ is a dummy variable taking value 1 if the firm has joint production in some other form like electricity and heat (otherwise variable takes value 2). Some auxiliary variables are also included in the regression. a_i is the firm specific variable for unobserved heterogeneity. λ_t is the common time effect for all firms describing the general business conditions.

$$\begin{aligned} lnEprice_{it} &= \alpha_i + \lambda_t + \beta_1 lnEprice_{it\text{-}1} + \beta_2 DJoint_{it} + \beta_3 DPUB_{it} \\ &\quad + \beta_4 lnMS_{it} + \beta_5 lnFUEL_{it} + \beta_6 lnSCALE_{it} + \beta_7 lnPRODS_{it} + \varepsilon_{it} \end{aligned} \tag{6)}$$

From the point view of consistent estimation the above fixed effects model (or its stochastic counterpart, random effects –model) is demanding. First, the inclusion of one period lagged endogenous variable in the panel models allowing for relevant price dynamics biases the LS estimates of model parameters. Second, to threat variables *DJoint*, *lnMS, lnSCALE,* and *lnPRODS* as exogenous variables is not warranted. The firm has a full control over its joint production possibilities and production scale. Thus the firm's pricing policy or the market price has effects on these variables. Similarly the causality between market share variables and heat tariff occurs most likely in both directions. If the firm has price setting power, then this has a feedback effect on its market share. Generally we can threat only the ownership and raw material unit cost variables as exogenous as their variability is most likely not affected by firm's pricing decisions.

Both these problems - biases in dynamic panel models and endogeneity – can be handled with instrumental variable estimation methods since technically question is of violation of so-called orthogonality conditions, i.e. non-correlation between the explanatory variables and error term. Instrumental variable estimation (IV) and related generalized methods of moments estimation (GMM) build on notion wherein we can find variables that are correlated with *lnEprice* variable but not with error term ε. Typically these are some variables outside the model and lagged values of right hand structural variables of the model.

Next we use following estimation strategy. First we estimate the equation in the pooled form, i.e. $\alpha_i = \alpha$ and λ_t reduces to single trend parameter. Next we estimate model with

IV method in panel form treating *DJoint*, *lnMS, lnSCALE,* and *lnPRODS* as endogenous. Finally we use GMM to estimate panel model in two forms: with and without the above endogeneity assumption. Note that GMM estimation is based on the one way random effects model where time trend is treated as a explanatory variable.

The estimation is done for three different data cases since the data contains three price variables: the energy tariff *(lnEprice)*, the total tariff for small houses $(lnEprice^{SM})$, and the total tariff for apartment houses and blocks $(lnEprice^{A})$. The total tariff has been selected because it includes all company specific factors which affect to the prices such geographical factors and population density. Similarly we have three different variables for market share of district heat among heat products, i.e. $lnMS,\ lnMS^{SM}$ *and* $lnMS^{A}$.

2.3. Price dynamics

A more elaborated theory pays attention to dynamics of price setting in the repeated game setting. The much discussed "Folk Theorem" says in this context that in a stationary oligopolistic situation in which players are impatient all the "cooperative" solutions of the one period game are the Nash equilibria of the overall non-cooperative game. Empirical testing is awesome task in this context and the results are indicative (e.g. Slade 1995, 1992, Gasmi et. al. 1992). Main econometric novelty in these papers is that firm's price setting is a function of others firms' lagged prices, i.e. $p_{it} = p_i(\sum_{j\neq i}^{N} p_{jt-1})$ corresponds to dynamic game setting.

Finally the cross-price dependence can be motivated by the fact that the firms' face a random demand and they have an option of information sharing concerning their private (cost) information (e.g. Clarke 1983, Vives 1984, 1990, Malueg & Tsutsui 1998, Zhu 2004). Basically results in this literature leads to formulations like

$$7) \qquad E[c_i \mid \sum_{j\neq i}^{N} c_j] = \mu + \frac{\rho}{1+\rho(N-1)} \sum_{j\neq i}^{N} (c_j - \mu)$$

where ρ is the correlation between firms' costs indicating that firm's unit costs (and price setting) depends on other firms' costs (prices).

2.3.1. Econometric specification

Assume that price process of individual firm can be expresses as the following non-homogenous first order difference equation

$$8)\qquad \Delta P_{it} = \alpha + \beta P_{i,t-1} + c\overline{P}_{j,t-1}$$

where $\overline{P}_{j,t-1} = \frac{1}{N-1}\sum_{j\neq i}^{N} P_{j,t-1}$ is the average price level of market excluding firm's i price at preceding period. The variable $\overline{P}_{j,t-1}$ measures the firm i's response to other firms' average pricing. We call it reference effect. In competitive environment it is expected that coefficient c takes zero or non-negative value. Coefficient β captures the dynamics and stability conditions of firm's own pricing process since the solution for Eq. (8) is

$$9)\qquad P_{it} = P_{i,0}(1+\beta)^{t} - \frac{1}{\beta}[\alpha + c\overline{P}_{j,t-1}].$$

If $-2 < \beta < 0$ and $[\alpha + c\overline{P}_{j,t-1}] > 0$ then the price process over the firms is stable and positive long run price level exists. When $\beta \approx -1$ we say that firm's pricing process does not have memory, and if $a, c > 0$, only the reference effect matters.

For view point of empirical testing the Eq. 8) is easily handled since $\overline{P}_{j,t-1}$ is independent and pre-determined of P_{it}. However $\Delta P_{it} = P_{it} - P_{i,t-1}$ may influence $\overline{P}_{j,t-1}$ causing endogeneity bias in OLS –estimation. The problem can be avoided with a modification typical to growth and convergence models, i.e.

$$10)\qquad \Delta P_{i,t}^{0} = \alpha + \beta P_{i,0} + c\overline{P}_{j,t-1} + \varepsilon_{it},$$

where $\Delta P_{it}^{0} = P_{it} - P_{i0}$ and ε_{it} is a IID error term. If the time points 0 and $t-1$ are far enough, then $\overline{P}_{j,t-1}$ is independent of ΔP_{it}^{0}, and $E[\varepsilon_{it}\overline{P}_{j,t-1}] = 0$.

If the price variability is high in the industry, the representative firm may find it difficult set its own price. If the firm has market power and face uncertainty in price setting then the firm overprices its output but the noise component (uncertainty) reduces the incentive for overpricing (i.e. $\partial P / \partial \sigma^2 < 0$, where σ^2 is the price variability). For more details see Appendix 2.

The price variability variable that excludes firm i's price is defined as

11) $$VAR[P_{t-1}]_j = \frac{1}{N-1}\sum_{j \neq i}^{N} (\overline{P}_{j,t-1} - P_{j,t-1})^2 .$$

Augmenting Eq. 10) with this gives

12) $$\Delta P_{i,t}^{0} = \alpha + \beta P_{i,0} + c\overline{P}_{j,t-1} + dVAR[P_{t-1}]_j + \varepsilon_{it},$$

It is argued that in non-competitive case $d < 0$: increased price variability reduced price changes as (monopoly) firm's gain from price increase comes less certain. The case $d \geq 0$ refers to competitive case wherein uncertainty is an exogenous cost factor.

Finally we augment the model with variable that measures the change in market power. Thus we add in model variable $\Delta PRODS_{i,t-1}^{0} = PRODS_{i,0} - PRODS_{i,t-1}$ to measure the change in market share of firm during the time period from 0 to $t-1$. We do not measure the change between periods 0 to t as the endogeneity between current period price dynamics and market share may be present. Now

13) $$\Delta P_{i,t}^{0} = \alpha + \beta P_{i,0} + c\overline{P}_{j,t-1} + dVAR[P_{t-1}]_j + \delta \Delta PRODS_{i,t-1}^{0} + \varepsilon_{it},$$

The same data set is used to estimate Eq. 13) as was used for the pricing level Eq. 6). However in this context we do not use any cost or scale etc. variables as the firm's relevant pricing decision information incorporates them in the market price variables $\overline{P}_{j,t-1}$ and $VAR[P_{t-1}]_j$, and in firm's own price and market share history variables $\Delta P_{i,t}^0$, $P_{i,0}$, and $PRODS_{i,t-1}^0$.

III. RESULTS

3.1. SCP model

Panel models

Table 2 and Tables A1-A2 in Appendix 1 give the results for different estimations with alternative data sets. We estimate the SCP-model for district heat price tariff, for total tariff for small houses (MS), and for apartment houses (A). Total tariff includes the taxes. Generally results are similar for different data sets. However results depend strongly on method used. We put more weight on $PANEL_{IV}^{2FE}$ and $PANEL_{GMM}^{E}$ estimation results than on two other method alternatives. Model diagnostics show that OLS is not adequate approach (large BP-test and Fixed Effects test values). $PANEL_{GMM}$ handles only the bias problem in panel setting but like OLS it is based on the restrictive assumption of exogeneity. In the following we analyze the results in details in context of different explanatory variables.

Year. Time trend or time effects (λ_t) are statistically significant in all estimations. Positive point estimate indicates that district heat tariffs have been increasing during the analyzed period.

$lnEprice_{-1}^{/MS/A}$. The significant lagged energy tariff implies that pricing is a dynamic phenomena and some adjustment is taking place in time. Point estimates range from *0.175* to *0.825* with average value of *0.426* revealing that it takes only some years to obtain steady state solution.

Table 2. OLS and panel model estimation results for district heat energy tariff in Finland *(lnEprice)*. Number of firms $N = 76$. Years 1997-2002, $T = 6$. (HCSE -corrected t-values in parenthesis).

	OLS	$PANEL_{IV}^{2FE}$	$PANEL_{GMM}$	$PANEL_{GMM}^{E}$
Constant	–9.461* (–1.65)	0.125 (0.188)	-20.02* (-2.72)	-21.83* (-2.24)
Year	0.005* (1.78)	-	0.011* (2.81)	0.012* (2.35)
$lnEprice_{-1}$	0.695* (16.22)	0.212* (3.31)	0.416* (5.00)	0.388* (4.12)
DJoint	0.029* (1.71)	0.041 (0.97)	0.058* (1.92)	-0.156 (-1.29)
DPUB	0.009 (0.80)	0.138* (2.74)	0.011 (0.77)	-0.130 (-1.18)
lnMS	-0.014 (-1.37)	0.094 (0.57)	-0.022 (-1.63)	-0.004 (-0.22)
lnFUEL	0.103* (5.04)	0.105* (2.64)	0.164* (6.54)	0.138* (3.58)
lnSCALE	-0.008 (-1.24)	0.380* (2.92)	-0.250* (-2.41)	0.031 (1.25)
lnPRODS	0.015* (2.95)	0.038 (0.82)	0.026* (3.50)	0.050* (2.80)
R^2	0.697	0.813	-	
DW / AC	2.16/-0.08	2.08/-0.05	-	
BP	125.71*	-	-	
FIXED EFFECTS	–	164.75*	-	
$Hausman/Sargan_{IV}$	–	2.73	92.81*	71.53*

*) significant at 10% critical level or below

$PANEL_{IV}^{2FE}$: 2-way fixed effect (2FE) panel data instrumental variable estimation (IV). Instruments: all exogenous variables (Year, DPUB, lnFUEL) and one year lagged values of lnFUEL and endogenous variables (lnMS, DJoint, lnSCALE, lnPRODS).

$PANEL_{GMM}$: Dynamic panel error components model GMM estimation. Instruments: see above

$PANEL_{GMM}^{E}$: Dynamic panel error components model GMM estimation with endogenous variables. Instruments: see above

DW/AC: Durbin-Watson test statistics and estimated residual 1[st] order autocorrelation

BP: Breusch-Pagan test for heteroskedasticity (H_0: no heteroskedesticity)

FIXED-EFFECT: Test for significance of firm specific and time effects (H_0: no fixed effects)

Hausman: Orthogonality test for valid instruments (H_0: $Corr(X_{IV}, \varepsilon)=0$)

$Sargan_{IV}$: Bhargava-Sargant test for instrument validity (H_0: moment conditions are valid)

$DJoint$. This joint production dummy variable is not well determined across the estimations. Only in 5 cases of 12 some statistical significance is obtained. The parameter point estimates are positive in relevant cases. We conclude that joint production does not decrease district heat tariffs.

$DPUB$. Public or state ownership does not decrease tariffs. Except for one estimation case this dummy variable takes non-significant value. Only in Table 1. with $PANEL_{IV}^{2FE}$ estimation we found positive and significant estimate.

$\ln MS^{/MS/A}$. The market share of district heat of total heat production enters significantly only once in all estimations. Combining this result with results for $DJoint$ variable indicates that horizontal heat product integration does not affect district heat pricing.

$lnFUEL$. The raw materials costs determine both in economic and statistical terms significantly the tariff level. Cost pricing rule is evident for heat power stations. However the input price elasticities are quite low ranking from 0.056 to 0.164. Long run estimates are at least twice larger. Measurement errors can cause partly the result but generally a low input elasticity is an implication of firm's pricing power: output prices do not reflect fully the marginal costs.

$lnSCALE$. Energy production theory and practice indicates that large plants are more effective, i.e. they can use scale effects to reduce marginal costs of output. However our results are somewhat mixed in this context. The scale parameter is statistically significant in 8 cases of 12 but we obtain both positive and negative estimate values. The sign depends directly on the exogeneity assumption. In estimations where scale of energy production is treated exogenous (*OLS* and $PANEL_{GMM}$) we obtain negative price effect. When treated endogenous, the scale of production increases heat tariffs. The former case refers to the competitive case. In the latter case higher prices are obtained with larger plants. If the large plant is a sole producer in district then the pricing refers to monopoly case. However a viable alternative explanation is that the large plants are less effective than the small ones. Typically these are public owned plants.

$lnPRODS$. The share of market that a firm has on the district heat production has a positive effect on the tariff, i.e. if the firm is a sole producer the energy price is higher than with many producers. This result is obtained with statistical significance in eleven cases of twelve (11/12). Parameter estimates range from *0.008* to *0.058* indicating that market power effect is not large. An increase in the firm's market power from 50% to 60% increases district heat tariff only in average about 0.60% in long run.

Semiparametric models

A more detailed picture of district heat product share effects on heat tariff are obtained with semiparametric regression estimate of Eq. 6). In this context estimation is done for different yearly observations with spline smoothing method. Now the product share effects on heat tariff are treated as a non-linear function $G(lnPRODS)$. In this way we obtain a nonlinear curve estimate describing the tariff effect at different product share levels.

$$\text{(6')} \quad lnEprice_{it} = \alpha + \beta_1 lnEprice_{it\text{-}1} + \beta_2 DJoint_{it} + \beta_3 DPUB_{it} + \beta_4 lnMS_{it} + \beta_5 lnFUEL_{it} + \beta_6 lnSCALE_{it} + G(lnPRODS_{it}) + \varepsilon_{it}$$

The smoothing-spline approach is the solution to the following programming problem given by Greene & Silverman (1994). Choose λ in such a way that a smooth estimate for $G(x_i)$ is obtained in minimizing the following penalized least-squares criterion

$$\text{15)} \quad \sum_{i=1}^{N} (y_i - \boldsymbol{X_i \beta} - G(x_i))^2 + \lambda \int_a^b (\frac{d^2 G(t)}{dt^2})^2 \, dt$$

where λ is a fixed constant, and $a \le x_1 \le .... \le x_n \le b$. The variables in the model are defined as $y_{it} = lnEprice_{it}$, $x_{it} = lnPRODS_{it}$, and $\boldsymbol{X}$'s are all other variables of the model that enter linearly. The model in Eq. 14) is estimated for each year separately.

Table 3. reports the results of linear part of the estimated semiparametric model. However the yearly tariff response curves to product share levels are more interesting. We find with Figures 3-8. that in years 1997, 2000, and 2001 a share above 30% has an increasing positive impact on tariff. Note that estimation was done with *ln* –transformed variables meaning that curve values below unit line correspond to negative tariff elasticity to output share, and values above unit line corresponds to positive point elasticity. In years 1998, 1999, and 2002 tariff response is decreasing with increasing share. However reliable decreasing response estimates are obtained only for year 1999 since the specification test by Hastie and Tibshirani (1990) does not reject the linear model alternative for years 1998 and 2002. Note also that linear OLS results for share variable *lnPRODS* are not significant (see Table 4) in these years. In years 1997 2000, and 2001 OLS gives positive and statistically significant coefficient estimates for share variable. The year 1999 is only exception form this rule. Semiparametric curve estimate is decreasing and OLS estimate for *lnPRODS* is negative. The result stem from the fact that district heat pricing went through some market changes in late 1990's. Household electricity markets in Finland were opened to the competition in the 1st November of 1998. The more detailed effects of this act were analyzed in companion paper by authors (Linden & Peltola-Ojala 2005).

In general semiparametric and cross-section OLS estimation results support the panel data estimation results. Increasing product market share has a positive effect on tariff price. Some yearly differences are found, especially for year 1999, and in some years like 1997 and 2000 a share less than 30% leads to price decreasing response. However these year specific cases do not rule out the general result that non-competitive price setting is dominant in the Finnish district heat markets.

Figure 3. Semiparametric estimate with 95% CI's of tariff response to district heat product share, 1997

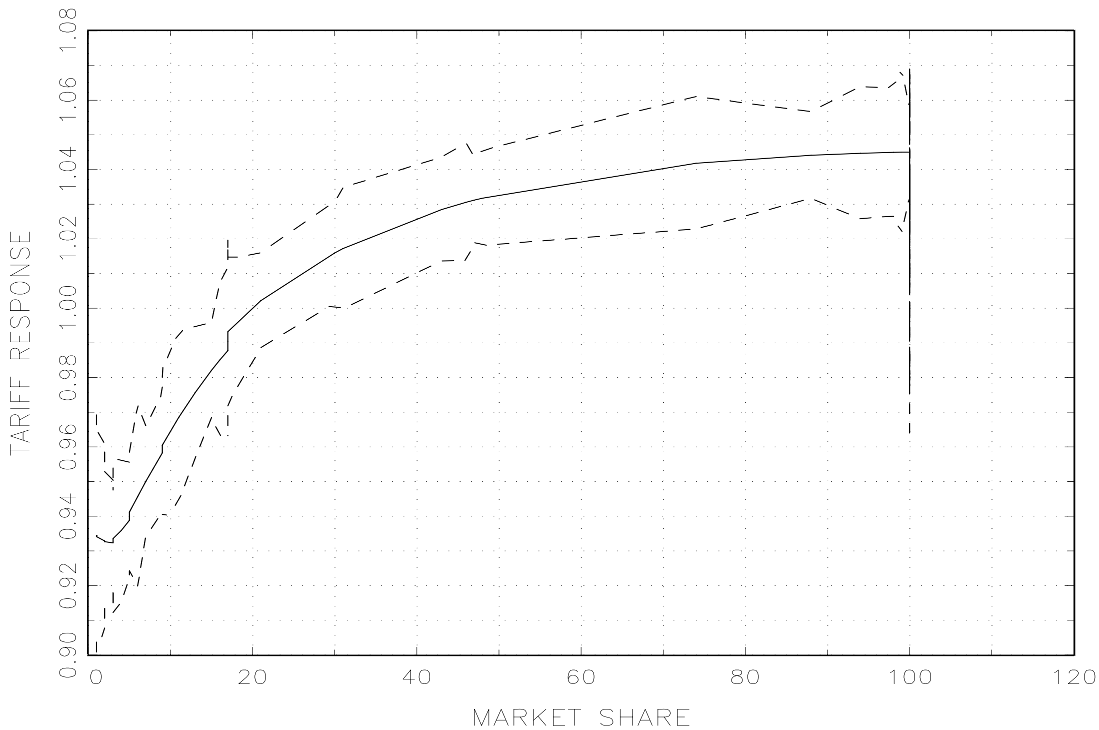

Figure 4. Semiparametric estimate with 95% CI's of tariff response to district heat product share, 1998

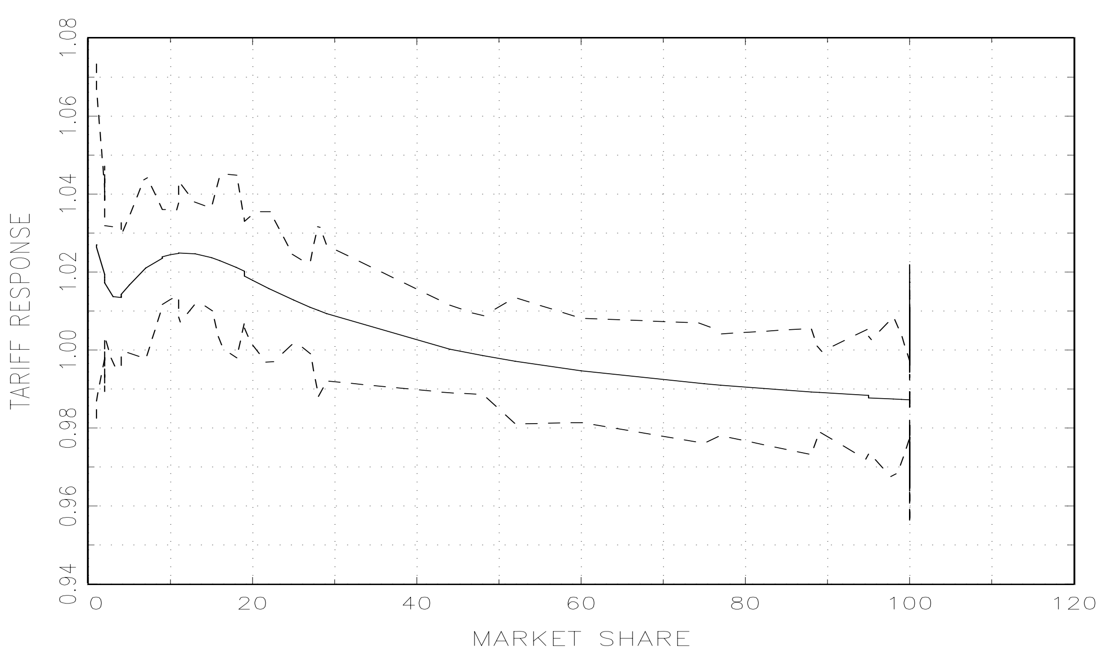

Figure 5. Semiparametric estimate with 95% CI's of tariff response to district heat product share, 1999

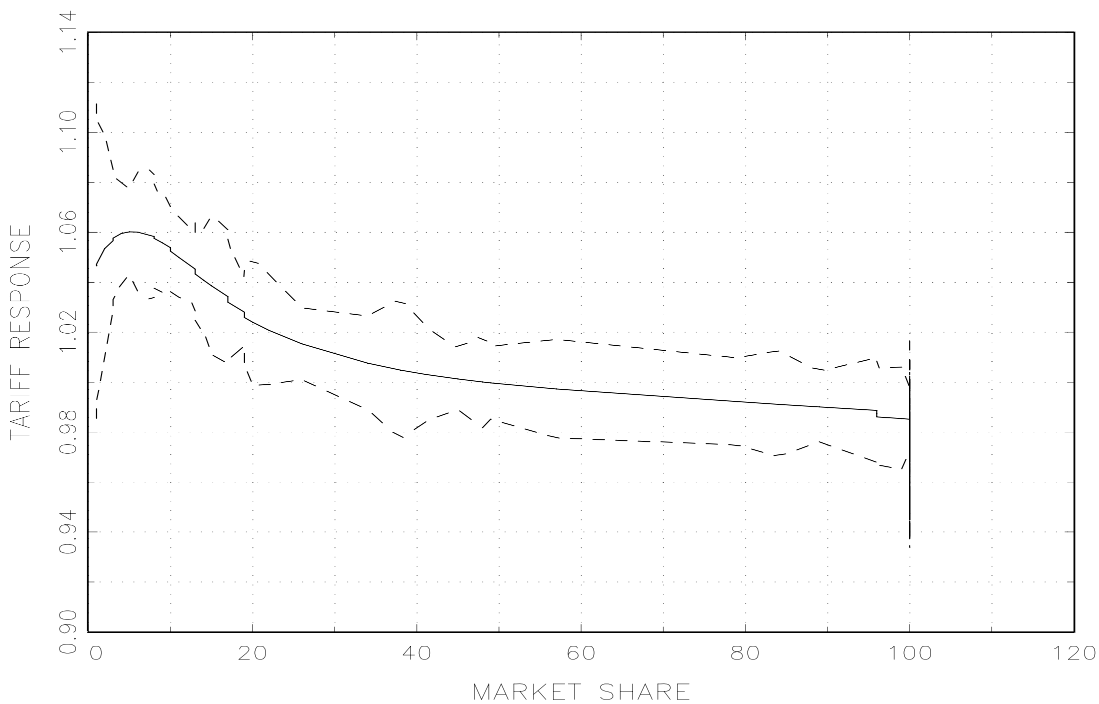

Figure 6. Semiparametric estimate with 95% CI's of tariff response to district heat product share, 2000.

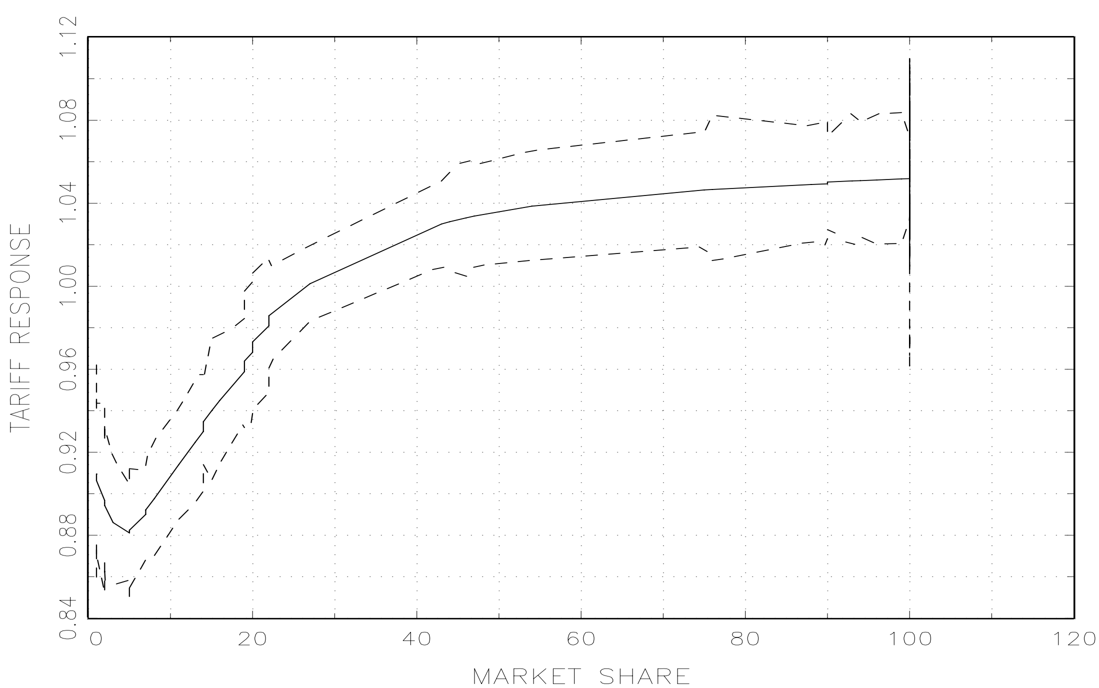

Figure 7. Semiparametric estimate with 95% CI's of tariff response to district heat product share, 2001.

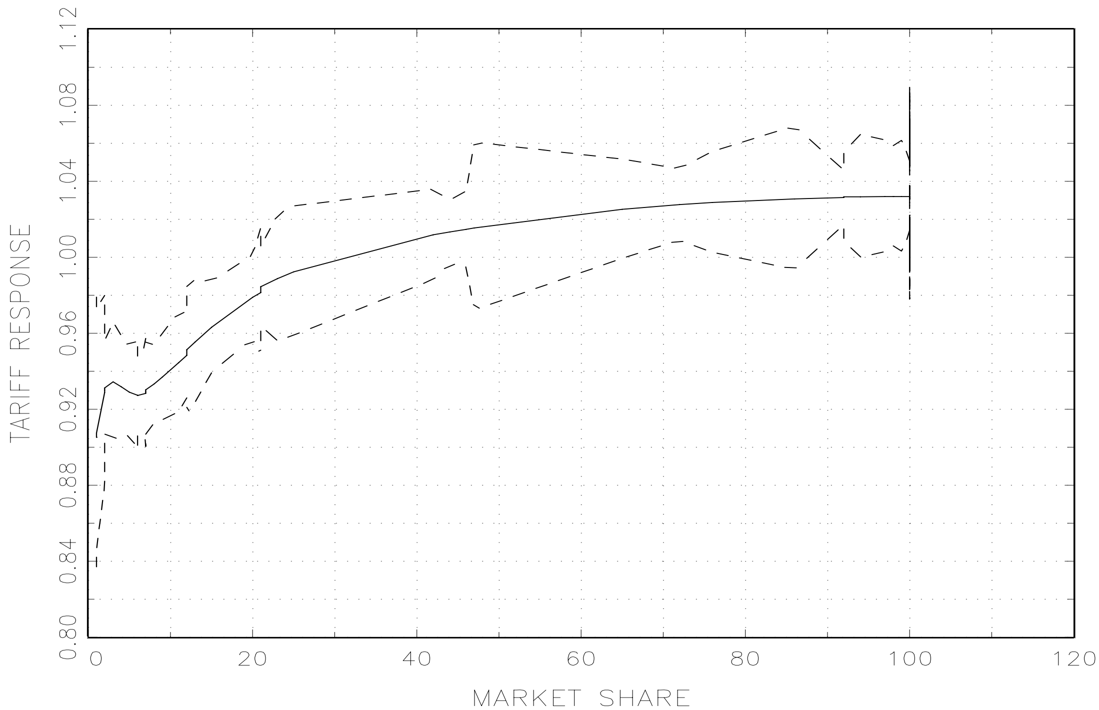

Figure 8. Semiparametric estimate with 95% CI's of tariff response to district heat product share, 2002.

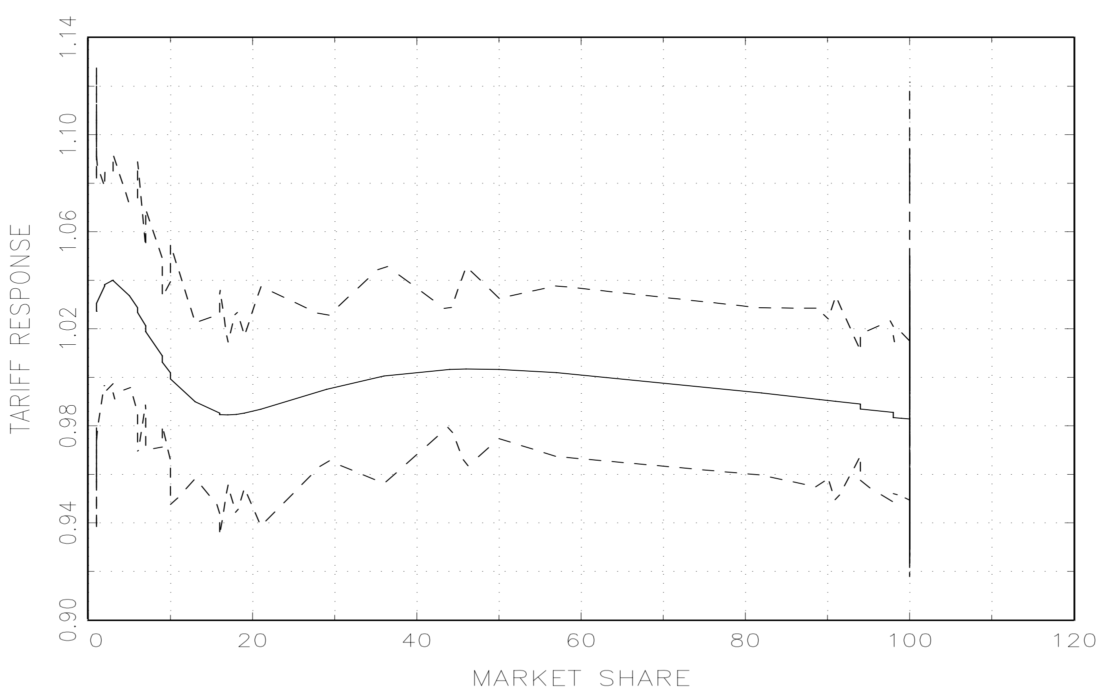

Table 3. Linear part of semiparametric cross section SCP -model (OLS estimation, HCSE – corrected t-values in parenthesis)

	1997	1998	1999	2000	2001	2002
Constant	0.516* (2.76)	0.417* (2.64)	0.249 (1.57)	0.964* (2.14)	0.037 (0.13)	1.14* (4.02)
$\ln Eprice_{-1}$	0.794* (15.04)	0.837* (15.91)	0.937* (16.26)	0.648* (4.86)	0.871* (9.31)	0.639* (8.57)
DJoint	–0.036 (–1.50)	0.086* (3.47)	0.047* (1.76)	-0.040 (-0.99)	0.048 (0.92)	0.040 (0.95)
DPUB	0. 003 (0.13)	0.038* (2.22)	-0.032* (-1.75	0.029 (0.91)	–0.047* (-1.99)	0.021 (0.82)
$\ln MS$	0.012 (1.16)	–0.023* (-2.51)	-0.007 (-0.63)	0.001 (0.08)	0.006 (0.47)	–0.010 (-0.56)
$\ln FUEL$	0.074* (2.98)	0.058* (1.83)	-0.038 (-1.50)	0.113* (4.31)	0.195* (5.31)	0.021 (0.46)
$\ln SCALE$	–0.009 (–0.06)	–0.006 (-0.52)	0.011 (0.99)	-0.018 (0.78)	–0.057* (-2.14)	0.006 (0.17)
R^2	0.871	0.891	0.811	0.689	0.816	0.557
H & T	4.95*	0.636	2.98*	5.58*	1.87	0.29
J & B	14.42*	16.73*	1498.11*	78.46*	12.33*	620.44*

H & T: Hastie & Tibshirani (1990) (asympt.) specification test for semiparametric model, H_0: model is linear (F-type test with $F(4,69)_{0.05}=2.53$)

J & B: Jarque & Bera (1980) (asympt.) test for residual normality ($\chi^2(2)_{0.05} = 5.91$)

Table 4. Cross section estimation of SCP –model (OLS estimation, HCSE –corrected t-values in parenthesis)

	1997	1998	1999	2000	2001	2002
Constant	0.432* (2.11)	0.429* (2.44)	0.307 (1.34)	0.949* (3.02)	–0.083 (-0.29)	1.18* (3.75)
$\ln Eprice_{-1}$	0.796* (13.25)	0.847* (15.13)	0.937* (12.90)	0.611* (6.41)	0.875* (9.76)	0.626* (6.29)
DJoint	–0.036 (–1.07)	0.085* (2.97)	0.036 (1.08)	-0.017 (-0.34)	0.052 (1.05)	0.034 (0.55)
DPUB	0. 002 (0.13)	0.040* (2.57)	-0.027 (-1.36)	0.021 (0.72)	–0.049* (-1.80)	0.023 (0.62)
$\ln MS$	0.013 (1.03)	–0.021* (-1.91)	-0.002 (-0.18)	-0.003 (-0.19)	0.005 (0.31)	–0.007 (-0.31)
$\ln FUEL$	0.074* (2.47)	0.054 (1.54)	-0.033 (-0.94)	0.117* (3.31)	0.192* (4.15)	0.026 (0.38)
$\ln SCALE$	–0.017 (–0.93)	–0.009 (-0.62)	0.011 (0.57)	-0.025 (-0.87)	–0.056* (-2.05)	0.004 (0.089)
$\ln PRODS$	0.024* (3.02)	–0.011 (-1.27)	-0.022* (-2.14)	0.038* (2.97)	0.032* (2.11)	–0.006 (-0.36)
R^2	0.860	0.887	0.796	0.649	0.807	0.530
B & J	21.88*	22.98*	1466.22*	70.41*	11.33*	635.21*

H & T: Hastie & Tibshirani (1990) (asympt.) specification test for semiparametric model, H_0: model is linear (F-type test with $F(4,69)_{0.05}=2.53$)

J & B: Jarque & Bera (1980) (asympt.) test for residual normality ($\chi^2(2)_{0.05} = 5.91$)

3.2. Price dynamics models

Tables 5a)-9a) give the price dynamics estimation results for all sample observations (N=76) over all possible yearly dynamics. Tables A5b) – A9b) in Appendix 1. give the results for sample that excludes all firms that have 100% market share in local district heat production (N=46). All the models were estimated with OLS –method.

Table 5a. OLS estimation of $\Delta P_{i,t}^{0} = \alpha + \beta P_{i,0} + c\overline{P}_{j,t-1} + dVAR[P_{t-1}]_j + \delta \Delta PRODS_{i,t-1}^{0} + \varepsilon_{it}$, where P_{i0} is year 1996 observations (N=76) (HCSE -corrected t-values in parenthesis).

	2002 – 1996	2001 – 1996	2000 – 1996	1999 – 1996	1998 – 1996
Constant	1490.82* (7.49)	2833.62* (8.69)	912.08* (2.73)	1288.42* (6.48)	1409.68* (6.72)
P_o	−1.01* (−7.78)	−1.12* (−6.76)	−0.51* (−3.34)	−0.76* (-7.85)	−0.77* (−7.22)
$\overline{P}_{t-1}$	−47.98* (−6.58)	−102.56* (−8.26)	−30.36* (−2.42)	-46.22* (−6.19)	−52.66* (−6.62)
$VAR[P_{t-1}]$	6.19 (0.77)	15.62* (1.66)	−22.90* (−2.04)	−4.48 (−0.87)	1.62 (0.31)
$\Delta PRODS_{i,t-1}^{0}$	0.01 (0.51)	−0.05 (−1.06)	0.02 (0.49)	−0.03 (−0.97)	0.01 (0.54)
R^2, *J & B*	0.533, 252.83*	0.548, 26.54*	0.181, 212.51*	0.491, 391.61*	0.466, 68.49*

J & B: Jarque & Bera (1980) (asympt.) test for residual normality ($\chi^2(2)_{0.05} = 5.91$)

Table 6a. OLS estimation of $\Delta P_{i,t}^{0} = \alpha + \beta P_{i,0} + c\overline{P}_{j,t-1} + dVAR[P_{t-1}]_j + \delta \Delta PRODS_{i,t-1}^{0} + \varepsilon_{it}$, where P_{i0} is year 1997 observations (N=76) (HCSE -corrected t-values in parenthesis).

	2002 – 1997	2001 – 1997	2000 – 1997	1999 – 1997
Constant	1400.91* (6.41)	2887.17* (7.05)	283.96 (1.20)	1950.41* (4.32)
P_0	−0.92* (−6.62)	−1.14* (−5.05)	−0.13 (−1.24)	-1.13* (−8.91)
$\overline{P}_{t-1}$	−44.88* (−5.72)	−104.28* (−6.87)	−7.17 (−0.81)	−71.10* (−7.25)
$VAR[P_{t-1}]$	4.83 (0.62)	14.64* (1.63)	−22.53* (-2.52)	0.16 (0.03)
$\Delta PRODS_{i,t-1}^{0}$	0.01 (0.61)	−0.05 (−1.11)	0.02 (0.49)	0.02 (0.41)
R^2, *J & B*	0.455, 270.58*	0.444, 26.32*	0.108, 9.98*	0.567, 696.49*

J & B: Jarque & Bera (1980) (asympt.) test for residual normality ($\chi^2(2)_{0.05} = 5.91$)

Table 7a. OLS estimation of $\Delta P_{i,t}^{0} = \alpha + \beta P_{i,0} + c\overline{P}_{j,t-1} + dVAR[P_{t-1}]_{j} + \delta\Delta PRODS_{i,t-1}^{0} + \varepsilon_{it}$, where P_{i0} is year 1998 observations (N=76) (HCSE -corrected t-values in parenthesis).

	2002 – 1998	2001 – 1998	2000 – 1998
Constant	1279.49 * (5.35)	2828.42 * (6.02)	−815.93 * (−2.74)
P_0	−0.83 * (−5.37)	−1.10 * (−4.57)	0.34 * (2.46)
$\overline{P}_{t-1}$	−40.59 * (−4.74)	−102.17 * (−5.86)	30.56 * (2.81)
$VAR[P_{t-1}]$	2.57 (0.32)	14.37 (1.53)	−1.19 (−0.14)
$\Delta PRODS_{i,t-1}^{0}$	0.01 (0.58)	−0.04 (−1.05)	−0.01 (−0.42)
R^2, $J \& B$	0.356, 325.22*	0.376, 25.78 *	0.105, 10.55*

J & B: Jarque & Bera (1980) (asympt.) test for residual normality ($\chi^2(2)_{0.05} = 5.91$)

Table 8a. OLS estimation of $\Delta P_{i,t}^{0} = \alpha + \beta P_{i,0} + c\overline{P}_{j,t-1} + dVAR[P_{t-1}]_{j} + \delta\Delta PRODS_{i,t-1}^{0} + \varepsilon_{it}$, where P_{i0} is year 1999 observations (N=76) (HCSE -corrected t-values in parenthesis).

	2002 – 1999	2001 – 1999
Constant	1279.47 * (6.51)	2713.73 * (8.64)
P_0	−0.79 * (−5.34)	−1.05 * (−6.10)
$\overline{P}_{t-1}$	−40.43 * (−5.63)	−98.01 * (−8.24)
$VAR[P_{t-1}]$	1.69 (0.21)	13.67 (1.42)
$\Delta PRODS_{t-1}^{0}$	0.01 (0.48)	−0.04 (−1.04)
R^2, $J \& B$	0.443, 401.81*	0.554, 26.53 *

J & B: Jarque & Bera (1980) (asympt.) test for residual normality ($\chi^2(2)_{0.05} = 5.91$)

Table 9a. OLS estimation of $\Delta P_{i,t}^{0} = \alpha + \beta P_{i,0} + c\overline{P}_{j,t-1} + dVAR[P_{t-1}]_j + \delta\Delta PRODS_{i,t-1}^{0} + \varepsilon_{it}$, where P_{i0} is year 2000 observations (N=76) (HCSE -corrected t-values in parenthesis).

	2002 – 2000
Constant	1372.82* (5.05)
P_0	–0.92* (–5.91)
$\overline{P}_{t-1}$	–44.08* (–4.74)
$VAR[P_{t-1}]$	5.21 (0.68)
$\Delta PRODS_{t-1}^{0}$	0.12 (0.57)
R^2, $J \& B$	0.335, 291.04*

J & B: Jarque & Bera (1980) (asympt.) test for residual normality ($\chi^2(2)_{0.05} = 5.91$)

P_o. Own price dynamics variable was statistically significant in all cases except for year 2000-1997 dynamics (Table 6a). The range of significant estimates was -1.14 - -0.51 with average of -0.93 excluding year 2000-1998 dynamics (Table 7a) that was positive with significant value of 0.34. The results indicate that own price memory effects are quite small in district heat pricing process. Pricing process is stable. No memory value $\beta = -1$ is inside the range of estimates but simple t-tests reject the hypothesis $\beta = -1$ (not reported). The non-significant value for price memory for year 2000-1997 (Table 5a) and positive value of β in year 2000-1998 (Table 7a) refer to the state induced competition incidence of the Finnish energy markets in late 1990's.

The results for sample excluding 100% market share firms (Tables A5b – A9b) are quite close to above ones. On the average the price dynamics has value of -0.95 but the range of estimates is somewhat wider (-1.46 - -0.55) than earlier. The induced competition effects for years 2000-1997 and 2000-1998 (Tables A6b-A7b) are clearly still observed.

$\overline{P}_{t-1}$. Market reference effects are statistically significant in all cases except for year 2000-1997 dynamics (Table 6a). The low variability of heat tariff price makes the estimates for parameter c to be quite large in absolute value. The range for significant estimates was -104.31 - -63.48 with average of -63.48 excluding year 2000-1998 dynamics (Table 7a) that was positive with significant value of 30.56. The results indicate that firms' react quite strongly to last period average market price. The impact is negative on firms' prices changes. This means that long run stable price level is found and a convergence toward it is established (see Eq. 9). Excluding the monopoly firms from the sample (Tables A5b-A9b) changes the results to form where reference effects is not on average so big absolutely (-35.77) and the range of estimates is more narrow than earlier (-53.64 - -18.54). The result indicate that if there are more than one firm on local district heat market the price reference effect is less than in the case were the monopoly firm are also present. Thus surprisingly the number of monopolies increases reference price effect.

$VAR[P_{t-1}]$. The price volatility measure has a statistically significant coefficient estimate only in four cases for whole sample. In years 2001 - 1996/7 volatility effect on heat tariff is positive but in years 2000 – 1996/7 it is negative. However in sample excluding the monopolies we get seven significant estimates for price volatility. They are all positive. Thus an increase in last period's price dispersion has an increasing tariff effect. This supports the competitive market hypothesis. Under marginal cost pricing increased uncertainty is a cost factor that increases output price.

$\Delta PRODS_t$. Change in firm's market share does not affect firm's price dynamics. The estimate of $\Delta PRODS_t$ coefficient is only in one case statistically significant and it has a negative sign (Table A6b: 1999-1997). The result is surprising since we obtained with panel models above a clear positive product output share level effect on heat prices. The result can be explained in many ways. The shares may have been quite stable during the analyzed period. Thus only the level effect is then obtained. $\Delta PRODS_t$ -variable may be redundant in the price dynamics model since price information contains all the relevant decisions information already.

3.3. Discussion

A quite different picture of district heat markets enters from the price dynamic model results compared to the results of SCP panel data model. However some similarities exist. Less competitive are the local district heat markets more price responsive are the firms. Thus monopoly firms response quite strongly to the average level of tariff in their pricing policy. However the response is stabilizing the price process. The result indicates that firms are sensitive to secure their market shares and positions and they react also to market information outside the local relevant district heat markets. The increase in price uncertainty has only effect when monopoly firms are excluded from sample. This and the fact that change in output shares do not effect price dynamics - but share levels correlate positively with price levels - indicates that market structure of the Finnish district heat markets is stable. Only in period of 2000-1999/8 when the induced competition was introduced by the state the market was disturbed (see also Linden & Peltola-Ojala 2005). Estimates of model of price dynamics break down during these years and output share effects were non-typical.

IV. CONCLUSION

The results obtained from different estimations above indicate clearly that competitive case is not the prevailing one in the Finnish district heat production in years 1996-2002. Although some econometric problems were not fully solved (i.e. the lack of suitable instrumental variables and non-normal model residuals) the results are not secondary. Point estimates for parameter for variables *lnMS, lnFUEL, lnSCALE,* and *lnPRODS* in pricing models together revealed features that are typical for non-competitive markets. Significant district heat price effect estimates were not found for district heat firm's market share in the *total heat product* markets (*lnMS*). The result supports our assumption that the markets of household heat products are defined independently. However, the firm's share effect of *district heat product* (*lnPRODS*) markets on district heat pricing was found to be positive and statistically significant. The result was strengthened with yearly based semiparametric model estimations. The firms' market shares in district heat production have an increasing or a non-decreasing positive effect on product tariff in all years, expect 1999. Plant scale effects (*lnSCALE*) were not unambiguously determined. Unit cost had positive, albeit small, effects on tariff prices. Tariff prices were not sensitive public ownership of plant but joint production had in some extension positive pricing effect.

These non-competitive market results were partly challenged by model estimates of price dynamics. The change in firm's market share ($\Delta PRODS$) does not determine tariff movements. In addition price uncertainty effects supported the competitive market structure alternative. However firm's pricing movements were very sensitive - albeit stability preserving - to changes in other firm's pricing decisions in preceding periods. The special circumstances of year 1999, deregulation of Finnish energy markets, were clearly observed. The stable and rigid Finnish district heat markets were disturbed extensively by induced competition. However the stability was quickly restored. Generally the results indicate that the Finnish district heat markets are non-competitive but entail price dynamics that are typical for stable oligopolistic markets.

Appendix 1.

Table A1. OLS and panel model estimation results for district heat energy total tariff in Finland for small houses $(lnEprice^{SM})$. Number of firms $N = 76$. Years 1997 – 2002, $T = 6$. (HCSE -corrected t-values in parenthesis).

	OLS	$PANEL^{2FE}_{IV}$	$PANEL_{GMM}$	$PANEL^{E}_{GMM}$
Constant	–5.321 (-1.61)	1.769* (4.63)	-7.520* (-2.31)	-6.401 (-1.43)
Year	0.003* (1.76)	-	0.005* (2.51)	0.004* (1.68)
$lnEprice^{SM}_{-1}$	0.825* (24.23)	0.242* (4.42)	0.562* (7.69)	0.562* (7.07)
DJoint	0.017 (1.14)	-0.112 (-0.32)	0.037* (1.92)	0.017 (0.29)
DPUB	-0.004 (-0.59)	0.017 (0.57)	-0.033 (-0.36)	-0.048 (-0.84)
$lnMS^{SM}$	-0.002 (-0.75)	-0.029 (-0.67)	-0.009* (-2.40)	-0.001 (-0.11)
lnFUEL	0.056* (3.89)	0.071* (3.05)	0.114* (6.97)	0.102* (4.87)
lnSCALE	0.003 (0.87)	0.195* (2.44)	-0.002 (-0.29)	0.004 (0.32)
lnPRODS	0.008* (2.51)	0.050* (1.76)	0.016* (3.64)	0.014 (1.61)
R^2	0.801	0.881	-	-
DW / AC	2.31/-0.15	2.12/-0.06	-	-
BP	154.20*	-	-	-
FIXED EFFECTS	–	84.70*	-	-
$Hausman/Sargan_{IV}$	–	2.83	112.65*	82.20*

*) significant at 10% critical level or below

$PANEL^{2FE}_{IV}$: 2-way fixed effect (2FE) panel data instrumental variable estimation (IV). Instruments: all exogenous variables (Year, DPUB, lnFUEL) and one year lagged values of lnFUEL and endogenous variables (lnMS, DJoint, lnSCALE, nPRODS).

$PANEL_{GMM}$: Dynamic panel error components model GMM estimation. Instruments: see above

$PANEL^{E}_{GMM}$: Dynamic panel error components model GMM estimation with endogenous variables. Instruments: see above

DW/AC: Durbin-Watson test statistics and estimated residual 1st order autocorrelation

BP: Breusch-Pagan test for heteroskedasticity (H_0: no heteroskedesticity)

FIXED-EFFECT: Test for significance of firm specific and time effects (H_0: no fixed effects)

Hausman: Orthogonality test for valid instruments (H_0: $Corr(X_{IV}, \varepsilon)=0$)

$Sargan_{IV}$: Bhargava-Sargant test for instrument validity (H_0: moment conditions are valid)

Table A2. OLS and panel model estimation results for district heat energy total tariff in Finland for apartment houses $(lnEprice^{A})$. Number of firms $N = 76$. Years 1997 – 2002, $T = 6$.(HCSE -corrected t-values in parenthesis).

	OLS	$PANEL_{IV}^{2FE}$	$PANEL_{GMM}$	$PANEL_{GMM}^{E}$
Constant	−9.699 * (-2.68)	2.140* (3.17)	-15.591* (-3.24)	-19.428* (-3.09)
Year	0.005* (2.90)	-	0.009* (3.53)	0.011* (3.26)
$lnEprice_{-1}^{A}$	0.702* (15.14)	0.175* (3.32)	0.439* (5.35)	0.409* (4.37)
DJoint	0.021* (1.71)	-0.026 (-0.66)	0.046* (2.26)	-0.003 (-0.04)
DPUB	-0.004 (-0.59)	0.041 (1.19)	0.002 (0.23)	0.021 (0.31)
$lnMS^{A}$	0.001 (0.13)	-0.142 (-0.95)	0.007 (0.74)	0.011 (0.78)
lnFUEL	0.074* (4.69)	0.086* (3.11)	0.116* (6.69)	0.089* (3.56)
lnSCALE	-0.012* (-2.51)	0.229* (2.55)	-0.029* (-3.79)	-0.018 (-0.94)
lnPRODS	0.013* (3.12)	0.058* (1.77)	0.020* (4.07)	0.019* (1.81)
R^2	0.725	0.838	-	-
DW / AC	2.26/-0.13	2.09/-0.06	-	-
BP	140.11*	-	-	-
FIXED EFFECTS	–	137.45*	-	-
$Hausman/Sargan_{IV}$	–	3.70	104.11*	80.66*

*) significant at 10% critical level or below

$PANEL_{IV}^{2FE}$: 2-way fixed effect (2FE) panel data instrumental variable estimation (IV).
Instruments: all exogenous variables (Year, DPUB, lnFUEL) and one year lagged values of lnFUEL and endogenous variables (lnMS, DJoint, lnSCALE, nPRODS).

$PANEL_{GMM}$: Dynamic panel error components model GMM estimation.
Instruments: see above

$PANEL_{GMM}^{E}$: Dynamic panel error components model GMM estimation with endogenous variables. Instruments: see above

DW/AC: Durbin-Watson test statistics and estimated residual 1st order autocorrelation

BP: Breusch-Pagan test for heteroskedasticity (H_0: no heteroskedesticity)

FIXED-EFFECT: Test for significance of firm specific and time effects (H_0: no fixed effects)

Hausman: Orthogonality test for valid instruments (H_0: $Corr(X_{IV},\varepsilon)=0$)

$Sargan_{IV}$: Bhargava-Sargant test for instrument validity (H_0: moment conditions are valid)

Table A5b. OLS estimation of $\Delta P_{i,t}^{0} = \alpha + \beta P_{i,0} + c\overline{P}_{j,t-1} + dVAR[P_{t-1}]_j + \delta \Delta PRODS_{i,t-1}^{0} + \varepsilon_{it}$, where P_{i0} is year 1996 observations (N=41) (HCSE -corrected t-values in parenthesis).

	2002 – 1996	2001 – 1996	2000 – 1996	1999 – 1996	1998 – 1996
Constant	854.73* (5.88)	1392.94* (6.22)	497.93* (1.86)	716.30* (6.03)	807.18* (5.44)
P_o	−1.10* (−6.18)	−1.03* (−4.69)	−0.55* (−2.48)	−0.82* (-7.23)	−0.76* (−5.33)
$\overline{P}_{t-1}$	−28.68* (−5.59)	−53.64* (−5.95)	−18.43* (−1.77)	-26.33* (−5.73)	−30.29* (−5.18)
$VAR[P_{t-1}]$	9.77* (1.67)	23.45* (2.39)	−0.97 (−0.09)	0.28 (0.08)	−2.44 (−0.46)
$\Delta PRODS_{i,t-1}^{0}$	0.01 (0.44)	−0.03 (−0.63)	−0.01 (−0.16)	−0.04 (−1.32)	−0.01 (-0.01)
R^2, *J & B*	0.547, 309.00*	0.537, 3.36	0.159, 170.57*	0.619, 25.69	0.484, 17.15*

J & B: Jarque & Bera (1980) (asympt.) test for residual normality ($\chi(2)_{0.05} = 5.91$)

Table A6b. OLS estimation of $\Delta P_{i,t}^{0} = \alpha + \beta P_{i,0} + c\overline{P}_{j,t-1} + dVAR[P_{t-1}]_j + \delta \Delta PRODS_{i,t-1}^{0} + \varepsilon_{it}$, where P_{i0} is year 1997 observations (N=41) (HCSE -corrected t-values in parenthesis).

	2002 – 1997	2001 – 1997	2000 – 1997	1999 – 1997
Constant	745.29* (4.40)	1292.44* (4.37)	−8.78 (−0.21)	1324.25* (7.94)
P_0	−0.92* (−4.41)	−0.92* (−3.23)	−0.01 (−0.03)	-1.46* (−8.93)
$\overline{P}_{t-1}$	−24.92* (−4.27)	−49.81* (−4.50)	1.91 (0.22)	−50.01* (−7.65)
$VAR[P_{t-1}]$	8.11 (1.32)	21.83* (2.59)	−9.50 (-1.36)	7.53* (2.11)
$\Delta PRODS_{i,t-1}^{0}$	0.01 (0.60)	−0.02 (−0.59)	0.01 (0.02)	−0.04* (−1.86)
R^2, *J & B*	0.387, 350.89*	0.403, 3.51	0.05, 0.762	0.567, 21.08*

J & B: Jarque & Bera (1980) (asympt.) test for residual normality ($\chi(2)_{0.05} = 5.91$)

Table A7b. OLS estimation of $\Delta P^0_{i,t} = \alpha + \beta P_{i,0} + c\overline{P}_{j,t-1} + dVAR[P_{t-1}]_j + \delta \Delta PRODS^0_{i,t-1} + \varepsilon_{it}$, where P_{i0} is year 1998 observations (N=41) (HCSE -corrected t-values in parenthesis).

	$2002-1998$	$2001-1998$	$2000-1998$
Constant	756.93* (4.57)	1325.92* (4.77)	−656.78* (−2.63)
P_0	−0.94* (−4.82)	−0.96* (−3.74)	0.41* (2.02)
$\overline{P}_{t-1}$	−25.30* (−4.04)	−51.17* (−4.87)	23.13* (2.49)
$VAR[P_{t-1}]$	8.18 (1.32)	22.37* (2.61)	9.04 (1.37)
$\Delta PRODS^0_{i,t-1}$	0.01 (0.55)	−0.03 (−0.62)	−0.04 (−1.27)
R^2, J & B	0.414, 375.71*	0.424, 3.65	0.213, 1.58

J & B: Jarque & Bera (1980) (asympt.) test for residual normality ($\chi(2)_{0.05} = 5.91$)

Table A8b. OLS estimation of $\Delta P^0_{i,t} = \alpha + \beta P_{i,0} + c\overline{P}_{j,t-1} + dVAR[P_{t-1}]_j + \delta \Delta PRODS^0_{i,t-1} + \varepsilon_{it}$, where P_{i0} is year 1999 observations (N=41) (HCSE -corrected t-values in parenthesis).

	$2002-1999$	$2001-1999$
Constant	730.16* (4.77)	1321.81* (5.95)
P_0	−0.88* (−4.43)	−0.94* (−3.83)
$\overline{P}_{t-1}$	−24.33* (−4.51)	−50.75* (−5.69)
$VAR[P_{t-1}]$	7.57 (1.20)	21.27* (2.13)
$\Delta PRODS^0_{t-1}$	0.01 (0.51)	−0.02 (−0.68)
R^2, J & B	0.404, 424.11*	0.527, 3.83

J & B: Jarque & Bera (1980) (asympt.) test for residual normality ($\chi(2)_{0.05} = 5.91$)

Table A9b. OLS estimation of $\Delta P_{i,t}^0 = \alpha + \beta P_{i,0} + c\overline{P}_{j,t-1} + dVAR[P_{t-1}]_j + \delta \Delta PRODS_{i,t-1}^0 + \varepsilon_{it}$, where P_{i0} is year 2000 observations (N=41) (HCSE -corrected t-values in parenthesis).

	2002 – 2000
Constant	935.75* (4.82)
P_0	−1.18* (−5.42)
$\overline{P}_{t-1}$	−31.35* (−4.70)
$VAR[P_{t-1}]$	10.40* (1.69)
$\Delta PRODS_{t-1}^0$	0.01 (0.39)
R^2, *J & B*	0.464, 291.86*

J & B: Jarque & Bera (1980) (asympt.) test for residual normality ($\chi(2)_{0.05} = 5.91$)

Appendix 2.

Assume that true market price is noisy in a following way

(A1) $$\tilde{v} = P + \varepsilon, \quad \text{where } \varepsilon \sim N(0, \sigma^2).$$

This means that at optimum the price is equal to expected market price and the quantity of supply to meet the demand must be set according to it

(A2) $$E[\tilde{v}] = P = P^* : \quad D_X(P^*) = X^*.$$

However the firm faces here a problem: the true or actual demand is determined by $D_X(\tilde{v})$. The firm must have some estimate of $D_X(\tilde{v})$ or $E[D_X(\tilde{v})]$ to obtain the correct value of P. Note that firm is able in average to set the price correctly. However risk of incorrect price setting is evident in any single event and the firm needs some "insurance" against it. The firm prefers small over-pricing to under-pricing as its target is to hedge against incorrect pricing.

The derive correct or the risk preserving price we take 2nd order Taylor approximation of $D_X(\tilde{v})$ around P^*

(A4) $$D_X(\tilde{v}) \approx D_X(P^*) + D_X{}'(P^*)(\tilde{v} - P) + \tfrac{1}{2} D_X{}''(P^*)(\tilde{v} - P)^2$$
$$= D_X(P^*) + D_X{}'(P^*)(\varepsilon) + \tfrac{1}{2} D_X{}''(P^*)(\varepsilon)^2.$$

Taking expectation of this one obtains

$$E[D_X(\tilde{v})] \approx D_X(P^*) + D_X{}'(P^*)E[\varepsilon] + \tfrac{1}{2}D_X{}''(P^*)E[\varepsilon]^2$$

(A5)

$$= D_X(P^*) + \tfrac{1}{2}D_X{}''(P^*)\sigma^2.$$

This means that a price P^0 exists with following properties if $D_X{}''(P^0) > 0$

$$D_X(P^0) = D_X(P^*) - \tfrac{1}{2}D_X{}''(P^0)\sigma^2 :$$

$$D_X(P^0) < D_X(P^*) \text{ with } P^0 > P^*$$

(A6)

and

$$P^0 = D_X^{-1}[D_X(P^*) - \tfrac{1}{2}D_X{}''(P^0)\sigma^2] \text{ with } \partial P^0/\partial\sigma^2 < 0.$$

The result gives the rationality for setting price above P^* as long as $\tfrac{1}{2}D_X{}''\sigma^2 > 0,$ i.e. the monotonic decreasing demand curve is convex for P^0. The economic explanation for over-pricing rests on the idea of minimizing the risk of incorrect price. The firm having price setting power sets a price markup on market price in face of uncertain demand. However the increasing noise in price variability decreases over-pricing. Figure 1 gives the graphical presentation of firm's pricing policy:

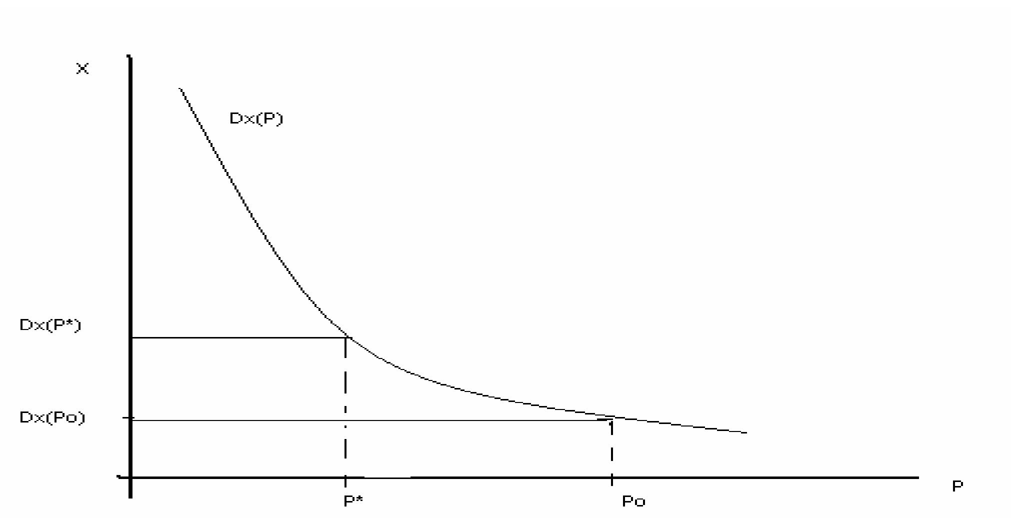

Figure A1. Firm's pricing under demand uncertainty:

$$\tfrac{1}{2}D_X^{''}(P^0) = D_X(P^*) - D_X(P^0) > 0\,.$$

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