PRICE RESPONSES TO CHANGES IN COSTS AND DEMAND

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STOCKHOLM SCHOOL OF ECONOMICS
EFI, THE ECONOMIC RESEARCH INSTITUTE

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Rickard Eriksson

Chapter 1

Introduction and Summary

Introduction and Summary

A large theoretical and empirical literature has studied the response in prices to changes in demands and costs. Three of the essays in this thesis are empirical studies of price-setting. The most important reason for studying the properties of price adjustments is the possible link between pricing and the business cycle. One family of models deals with price rigidities. If prices adjust slowly to changes in costs or demand, it will magnify fluctuations in output. A second family of models examines price changes due to changes in intensity of competition. Many of these models build on the idea that changes in demand affect the possibility of maintaining implicit collusion. Cyclical changes in the intensity of competition may cause a cyclical pricing pattern that magnifies fluctuations in output. Another line of research models the effects of liquidity constraints on the business cycle. One possible effect is that prices in markets where consumers have switching costs may increase, if the firms are hit by increased liquidity constraints. As increased liquidity constraints are common in recessions, prices will get a counter-cyclical tendency, which magnifies fluctuations in output. Another reason for studying price-setting patterns is that they can give an indication of the form of interaction between competing firms, which are of importance for competition policy. The fourth essay models sex discrimination formally. It shows that it is possible that sex discrimination in the labor market might be due to self-fulfilling sex stereotypes on the distribution of time out for child care between men and women. Both an even and an uneven distribution of time out for child care are possible equilibria in the model. The four essays can be read separately

Essay 1 Price Responses to Seasonal Demand Changes in the Swedish Gasoline Market

A common idea in a number of papers on cyclical pricing is that implicit price collusion may be affected by changes in demand. In these models, tacit collusion is modeled as a game where agents balance gains from deviating from the collusive price, thereby gaining a short-run profit, against the gains from maintaining collusion in future periods. If demand fluctuates, the gains from deviating is high in periods of high demand; collusion may still be sustainable, however, if the collusive price is allowed to vary with demand. A lower price in high demand states reduces the gains from deviating, since it reduces the gain from each unit sold in these demand states. Hence, in order to equalize the profit from deviating and the profit from sticking to the implicit agreement, the price must move in the opposite direction to demand. Changes in demand can, naturally, be correlated with other factors affecting the price, such as cost changes. Thus, prices do not necessarily fall when demand increases in the data. According to the theory the, however, increase in demand should add a tendency to lower prices in order to make implicit collusion sustainable. The basic idea that changes in demand should affect prices if firms are engaged in implicit collusion has been employed in a number of papers, with different models for different assumptions on the pattern of demand changes. The demand for gasoline in Sweden follows a seasonal cycle, with demand being 42 percent higher in July than in January. Haltiwanger and Harrington provide a theory for implicit collusion over deterministic cycles, such as the seasonal cycle. Borenstein and Shepard (1996) tested Haltiwanger and Harrington's model on the American gasoline retail market and found support for the theory. In this essay the model is tested for Swedish data, but no support for this theory is found. It is also investigated whether the effects on margins of the demand fluctuations induced by tax increases are compatible with theories of implicit collusion, but this is found not to be the case.

Essay 2 Price Adjustments by a Gasoline Retail Chain

Essay 2 is joint with Marcus Asplund and Richard Friberg. Stickiness of prices is an important building block in many business cycle models. This has spurred an empirical literature on price adjustments. Different types of price rigidities have different policy implications. Price setting can be state dependent, e.g. prices are adjusted when costs have changed by at least some minimal amount since the last price adjustment, or time-dependent e.g. adjusted once a week or once a year. Another issue is whether prices are equally rigid downwards as upwards. The second essay examines price responses in the Swedish gasoline retail market to changes in the Rotterdam spot price of gasoline, exchange rates and taxes. The main results are that cost changes are not fully passed through in the short run, but gradually moves towards the long-run equilibrium. Prices are stickier downwards than upwards. There is a minimum absolute size of price changes. Only very limited evidence of time-dependent price setting is found.

Essay 3 Prices, Margins and Liquidity Constraints: Swedish Newspapers 1990-1996

Essay 3 is written with Marcus Asplund and Niklas Strand. Chevalier and Scharfstein (1996) provide a model where consumer switching costs in combination with liquidity constraints give rise to a counter-cyclical tendency in prices. Customer stocks can be viewed as an investment, when consumers have switching costs when changing suppliers. Firms can exploit captured customers by setting a high price to raise short-run profits. However, a high price will induce consumers to search for alternatives, and customers once lost are costly to win back. Liquidity constraints, for instance in recessions, may force firms to sacrifice long-run profits for short-run gains. In this case, firms may have to cut back on investments in customer stocks by raising prices. Using firm level data from the Swedish daily newspaper industry, we test the effects of liquidity constraints on prices in markets with consumer switching costs. The newspaper industry is of particular interest, since firms set prices in two markets, the subscription market, where switching costs are high, and the advertising market, where switching costs are low. With accounting data from newspaper firms we can, by solvency, broadly categorize them as being more or less liquidity constrained. When Sweden enters a recession at the beginning of the nineties, we find a relative increase in subscription prices and margins for liquidity constrained firms. This is not the case for advertising prices, however. The results support the theory.

Essay 4 Statistical Discrimination and Sex Stereotypes in the Labor Market

Time out for child care is unevenly distributed between the sexes. Parental leave benefits are usually exclusively given to the mother or distributed to both parents according to their choice. In Sweden, one month is reserved for each parent, however. One reason for this is to give the child better contact with both parents, but increased equality between the sexes in the labor market has also been put forward as an argument. This argument implicitly rests on the idea that sex stereotypes create sex discrimination, and that sex discrimination affects the distribution of time out for child care between the sexes. Essay 4 investigates if the uneven distribution of time out for child care can be explained by self-fulfilling sex stereotypes. It provides a model of distribution of time out for child care based on statistical discrimination and human capital investments. The model has three equilibria. In one equilibrium, time out for child care is evenly distributed between the sexes. In the second equilibrium, there is full specialization. The third equilibrium is an intermediate case, where time out for child care is unevenly distributed without full specialization There are no differences in ability or variance of ability between the sexes, the only differences between the equilibria are the self-fulfilling expectations of firms and workers.

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Chapter 2

Essay I: Price Responses to Seasonal Demand Changes in the Swedish Gasoline Market

Price Responses to Seasonal Demand Changes in the Swedish Gasoline Market*

by

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Abstract

The demand for gasoline in Sweden follows a seasonal cycle. The paper investigates the response in prices and profits over the cycle. In contrast to what has been found for the gasoline market in the United States, we find no support for seasonal price changes compatible with the theories for cyclical variations of intensity of competition. We also investigate whether the effects on margins of the demand fluctuations induced by tax increases are compatible with these theories and find this not to be the case. Some possible explanations for this difference between Sweden and the United States are discussed.

Key words: Seasonal cycles; gasoline market.

JEL classification: E320; L130; L710

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1 Introduction

A large theoretical and empirical literature has studied the relationship between cyclical demand and prices or margins. The most important reason for studying the cyclical properties of prices is the possible link between cyclical pricing and the business cycle. Prices will tend to be counter-cyclical if increasing volumes have a reducing effect on prices. This will then magnify fluctuations in output. Several theoretical motivations for cyclical changes in prices have been proposed. One family of models has studied the relation between cyclical demand changes and the intensity of competition. In this paper, we test one of these models, Haltiwanger and Harrington's (1991) model for pricing over a deterministic cycle. The Swedish gasoline retail market is well suited for testing the theory, since variations in demand follow a deterministic seasonal cycle over the year. Borenstein and Shepard (1996) tested Haltiwanger and Harrington's model on the American gasoline retail market and found support for the theory. For Swedish data we find no support for this theory. One reason might be that the market structure differs between Sweden and the United States. A question of interest is therefore whether the difference in market structure affects the cyclical properties of prices. Although this paper only tests one specific model for a seasonal cycle, it can also be viewed as a test of the more general idea that changes in demand may affect the intensity of competition. In this way, the test is also relevant for the question of pricing over the business cycle.

The outline of the paper is as follows. Section 2 gives the background to and a summary of the model tested. Some empirical research is also reviewed. Section 3 describes and explores the data. In section 4, the theory is tested. We find no support for the model in this section. The price response to large tax increases is discussed in section 5, and the theory is not supported by the results in this section either. Section 6 discusses the results. It is found that the difference in market structure is a plausible explanation for the different results for the Swedish and the American market.

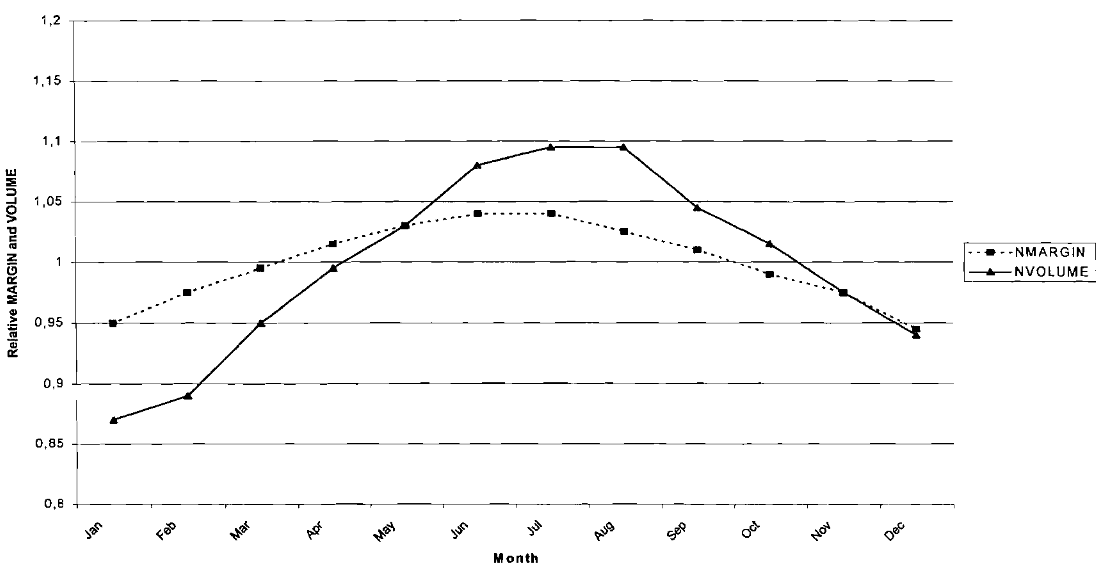
2 Literature review

Tacit collusion is often modeled as the outcome of a game where agents balance the gains from deviating from the collusive price, thereby gaining a short run profit, against the gains from maintaining collusion in future periods. If demand fluctuates, the gain from deviating is high in periods of high demand. Collusion may still be sustainable if the collusive price is allowed to vary with demand. A lower price in high demand states reduces the gains from deviating, since it reduces the gain from each unit sold in these states. This gives rise to a cyclical price pattern over the business or seasonal cycle. The first paper that formally models this idea is Rotemberg and Saloner (1986). They provide a formal model of collusion over varying demand states, which predicts that an increase in current demand over future demand will have a negative effect on prices. The demand states are assumed to be identically and independently distributed over time, so that expected future demand is equal in all periods. A realization of an unusually high demand state makes it more profitable to deviate, given the price of the competitors, since current demand is higher than normal, but expected future demand remains unchanged. Hence, in order to equalize the profit from deviating and the profit from sticking to the implicit agreement, the price must move in the opposite direction to demand.

Other papers have employed Rotemberg and Saloner's basic idea, but with different assumptions for the evolution of demand. Bagwell and Staiger (1997) assume changes in demand to be serially correlated, which leads to fast and slow growth phases of demand. Demand for gasoline in Sweden follows a similar pattern for all years in the period studied. This is not consistent with the assumptions in Rotemberg and Saloner or Bagwell and Staiger. Haltiwanger and Harrington provide a model for a deterministic demand cycle, which is more appropriate for the Swedish gasoline market. Their model is quite different from that of Rotemberg and Saloner, but both models build on the idea that the collusive price depends on the relation between current and future demand. Demand is assumed to rise in each period until reaching the peak level, then it falls in each period until it reaches the lowest level of demand. This is the only restriction Haltiwanger and Harrington impose on the demand cycle; there are no other restrictions on the speed or duration of the changes in demand over the cycle. The model gives different predictions for different discount factors. Firms will collude at the monopoly price level in all periods if the discount factor is sufficiently high and price at the marginal cost level in all periods if the discount factor is sufficiently low. The most interesting analysis is in intermediate case. Firms will then collude, but the sustainable collusive price will vary over the cycle, with prices below the monopoly price in at least one period and above the marginal cost in at least one period.

Haltiwanger and Harrington's model provide two testable predictions. If the discount factor is high enough, so that prices exceed marginal costs, profits will weakly lead the cycle. The second prediction is that, controlling for the level of demand, prices will be higher when demand is increasing than when it is decreasing.

Diagram 1: Developement of normalized margin and volume in the United States



There is a relatively large literature studying price setting on the gasoline market. Two of the papers are of particular interest. Borenstein and Shepard (1996) study the response in prices to seasonal demand changes in the United States, i. e. the same issue as we study for Sweden. They have a panel consisting of 43 markets over 6 years. The chain between the international market and the gasoline retail market consists of several links in the United States. Cost changes are gradually passed through from the international to the regional and local markets and hence, expected marginal cost changes must be included in the estimation of the margin. A large fraction of gasoline sales is unbranded and sold by independent retailers.

Diagram 1 shows the evolution of normalized margin and normalized volume in the United States. (Source: Borenstein and Shepard 1996.) The effects of different short-run dynamics in e. g. the pass-through of costs to prices are removed in the measure of margin. The pattern is consistent with the predictions offered by Haltiwanger and Harrington. Volumes lead margins. The margin is higher in periods when demand is expected to increase than when it is expected to decrease, for periods with approximately the same demand. Borenstein and Shepard's econometric investigation also supports the theory.

The short-run price dynamics of the Swedish gasoline retail market has been studied by Asplund, Eriksson and Friberg (2000). The main results were that cost changes were gradually passed through and that the price adjustment to cost increases and cost decreases was asymmetric in the short run, but symmetric in the long run. However, the price response to seasonal demand changes was not studied.

3 Data description

We study the seasonal pattern of margins for the Swedish gasoline chains for the period 1980-1996. Demand for gasoline increases in the spring, peaks in the summer, declines in the fall and passes a trough in the winter. Sales are, on average, 42% higher in July than in January. The seven largest firms have a total market share of 95%, but no single firm has a market share exceeding 25%. The total value of gasoline sales is about 3% of GDP in Sweden. Taxes account for a large fraction of the price of gasoline. During the period studied taxes constitute on average about 55% of the retail price of gasoline, varying from 40% to 71%.

The gasoline price used in this paper is the VAT-excluded list price for premium leaded gasoline. The price is usually the same for all firms, and a large dispersion in prices is very uncommon. We use the price for one of the firms, Shell, referred to as the firm below. The results are not sensitive to what firm is chosen. Gasoline is, at least physically, a relatively undifferentiated good. Consumers have a low inventory capacity, which puts a limit on the extent to which sales can increase when the price is unusually low, e. g. before a tax increase or during a price war.

The chain between the international spot market price and the retail price is simpler in Sweden than in the United States. In Sweden, the retail price is directly linked to the Rotterdam spot market price. Some firms buy gasoline at the Rotterdam spot market, whereas firms with their own gasoline production use the Rotterdam spot price as the transfer price. Firms either own the gasoline retail stations or set the price for their franchisees.

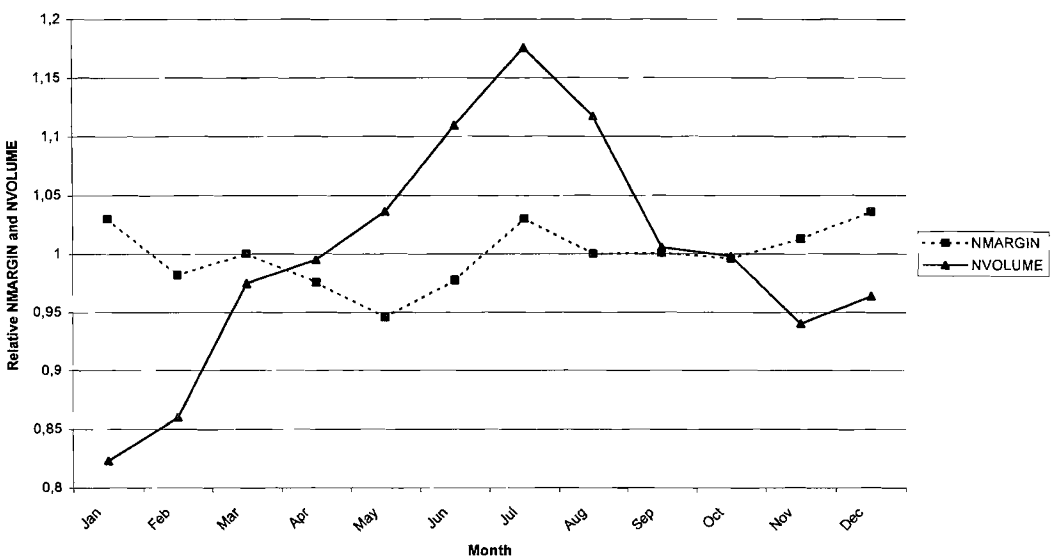
The margin, *MARGIN*, is defined as the retail price, *RP*, minus the per liter tax, *TAX*, and the cost for buying gasoline at the Rotterdam spot market, *MC*. (Note that *MARGIN* is measured in SEK*100, not as the percent mark-up.) The Rotterdam gasoline price is denoted in USD, and is multiplied by the SEK/USD exchange rate in order to obtain *MC* in SEK*100. The retail price, *RP*, the Rotterdam gasoline price and the SEK/USD exchange rate are available on a daily basis. *MONTHVOL* and *YEARVOL* are the monthly and yearly volumes sold. We use price and cost data as of the fifteenth day of each month in the regressions, since we only have access to monthly data on quantities. Table 1 displays the descriptive statistics for the variables.

Table 1: Descriptive statistics

	Mean	*Std dev*	*Min*	*Max*	*N*
MARGIN	87.30	29.40	7.32	147.76	204
MC	119.95	34.65	66.98	197.30	204
TAX	281.56	97.46	139.2	442	204
RP	488.81	105.69	276	672.8	204
MONTHVOL	448.45	60.16	303	561	204
YEARVOL	5381.18	427.85	4679	5910	204
SEK/USD	6.68	1.17	4.12	9.66	204

MARGIN, *MC*, *TAX* and *RP* are measured in nominal SEK*100 and *MONTHVOL* and *YEARVOL* in 1000 m^3. *TAX* and sold volumes are reported in the annual reports from the Swedish Petroleum Institute. Platt's is the source for *MC*, a firm collecting prices in the oil market. There is a positive trend in *MARGIN*, *TAX* and *RP* for the period studied, which is expected since these variables are nominal, but there is no trend in *MC* due to falling real prices on oil. The variance in *YEARVOL* is rather low. The highest yearly volumes are from the end of the eighties. Most of the variation in *MONTHVOL* is explained by seasonal variation. There is no systematic seasonal variation in the use of rebates, but there is some seasonal variation in transportation costs to the northern parts of Sweden, which is not accounted for in the measure of *MC*. According to the firm, this variation is not passed through to the prices and hence not to *MARGIN*. These cost changes is below 1 SEK*100. The results in the next section are not sensitive to this

Diagram 2: Developement of average normalized margin and volume



measurement error and remain almost unchanged if *MARGIN* is decreased by 1 SEK*100 during the winter months.

Diagram 2 shows the development of average normalized margin and average *MONTHVOL*. Both *MONTHVOL* and profits (measured as *MARGIN*MONTHVOL*) peak in July, which is consistent with Haltiwanger and Harrington's prediction that profits weakly lead demand. However, the prediction that, for a given demand, prices should be higher, when demand is expected to increase finds no support in the diagram. If anything, the opposite seems to be the case. The development of normalized margin is quite different from Borenstein and Shepard's data for the gasoline market in the United States. The most important difference is that, for a given level of demand, normalized margin tends to be high when demand is expected to increase. Demand in, for example, May and September is approximately of the same magnitude, but the margin is much lower in May.

4 Econometric Analysis

Already by looking at Diagram 1, one may suspect that it will be hard to find support for Haltiwanger and Harrington's model in the data. In this section we will present the results from some econometric specifications which formally test the theory. None of them support the theory, however.

Part I

The first regressions follow Borenstein and Shepard's approach as closely as possible in order to facilitate comparisons with their results for the gasoline market in the United States.

RP, *MC* and *TAX* are cointegrated. The null hypothesis of unit root cannot be rejected at the 10 percent level for any of the variables. A Johansen cointegration test including a constant and a linear trend rejects the null hypothesis of no cointegration at the 5 percent level. The inclusion of the constant and the linear trend reduces the Akaike information criterion, hence they are included in the cointegrating relationship.

The cointegrating relationship:

$$RP = \alpha + \beta_1 TIME + \beta_2 MC + \beta_3 TAX \quad (1)$$

We specify an error correction model in which deviations from the cointegration relationship as well as a number of variables important for the short run dynamics are included. Furthermore, we distinguish between positive and negative changes in MC, since the response in prices in previous studies of gasoline has often been shown to be asymmetric. See Bacon (1991) for British data, Borenstein, Shepard and Cameron (1997) for American data and Asplund, Eriksson and Friberg (1999) for Swedish data. The last four variables are the error correction term. The constant in the cointegrating relationship is implicitly included in α.

$$RP - RP_{-1} = \alpha + \beta_1 MONTHVOL + \beta_2 \mathrm{E}(MONTHVOL) + \beta_3(\Delta MC|\Delta MC>0) + \beta_4(\Delta MC|\Delta MC>0)_{-1} + \beta_5(\Delta MC|\Delta MC<0) + \beta_6(\Delta MC|\Delta MC<0)_{-1} + \beta_7 \Delta TAX + \beta_8 \Delta TAX_{-1} + \beta_9(\Delta RP|\Delta RP>0)_{-1} + \beta_{10}(\Delta RP|\Delta RP<0)_{-1} + \beta_{11} TIME + \beta_{12} RP_{-1} + \beta_{13} MC_{-1} + \beta_{14} TAX_{-1} \tag{2}$$

By rewriting equation 2 we get an expression for *MARGIN*, which is the variable of interest.

$$MARGIN = \alpha + \beta_1 MONTHVOL + \beta_2 \mathrm{E}(MONTHVOL) + \beta_3(\Delta MC|\Delta MC>0) + \beta_4(\Delta MC|\Delta MC>0)_{-1} + \beta_5(\Delta MC|\Delta MC<0) + \beta_6(\Delta MC|\Delta MC<0)_{-1} + \beta_7 \Delta TAX + \beta_8 \Delta TAX_{-1} + \beta_9(\Delta RP|\Delta RP>0)_{-1} + \beta_{10}(\Delta RP|\Delta RP<0)_{-1} + \beta_{11} TIME + (\beta_{12}+1) RP_{-1} + (\beta_{13}-1) MC_{-1} + (\beta_{14}-1) TAX_{-1} \tag{3}$$

The only forward-looking variables in the regression are the expected changes in volume of sales. The marginal cost is a random walk, so there is no need to include any measure of expected change in marginal cost.

Expected instead of actual volume of sales is used to circumvent the potential simultaneity problem caused by both *MARGIN* and actual sales being dependent on the price. This may not be a very serious problem, however, since short-run demand elasticity is known to be very low.

Table 2: Estimation of expected volumes sold

Variable	
CONSTANT	-74.6***
	(20.1)
FEB	11.8
	(7.29)
MAR	60.0***
	(7.31)
APR	74.5***
	(7.23)
MAY	91.1***
	(7.27)
JUN	119.6***
	(7.31)
JUL	154.4***
	(7.23512)
AUG	126.5***
	(7.27)
SEP	72.6***
	(7.31)
OCT	76.5***
	(7.23)
NOV	43.9***
	(7.29)
DEC	55.6***
	(7.43)
GDP	0.000329***
	(0.00000142)
$TAXINC_{t-1}$	7.14
	(8.92)
TAXINC	-38.8***
	(8.93)
$TAXINC_{t+1}$	68.0***
	(8.94)
Adj R^2	0.8790
D-W	2.10
N of obs	203

Variables starred *, ** and *** indicate significance at the 10% 5% and 1% level, respectively. Standard errors in parenthesis.

The regression for estimating the expected volumes sold, E(*MONTHVOL*), is shown in Table 2. An R^2 of 0.86 indicates that most of the variation in volumes sold is due to changes in demand. Price changes in response to tax changes are known in advance and are often of a larger magnitude than other price changes. The tax changes therefore affect demand in the period they occur and in the periods immediately before and after the tax change. (See the next section for further discussion.) A dummy for tax changes larger than 15 SEK*100 is included in the regression. Smaller tax changes do not affect demand, see Table 7 in section 5. Including the magnitude of tax changes or including all tax changes would be an alternative. The results are not sensitive to which of these specifications is chosen. A linear trend was included in preliminary regressions, but turned out to be insignificant and did not affect any results.

Table 3: Dependent variable *MARGIN*

	OLS	*2SLS*
CONSTANT	50.3***	48.4***
	(10.8)	(10.7)
TIME	0.0112***	0.0109***
	(0.00257)	(0.00258)
RP_{-1}	0.654***	0.661***
	(0.0601)	(0.0599)
MC_{-1}	-0.699***	-0.703***
	(0.0594)	(0.0598)
TAX_{-1}	-0.779***	-0.782***
	(0.0450)	(0.0449)
MONTHVOL	-0.0162	-0.0115
	(0.0160)	(0.0181)
E*(MONTHVOL)*	-0.0172	-0.0199
	(0.0167)	(0.0175)
$(\Delta MC \mid \Delta MC>0)$	-0.315***	-0.308***
	(0.101)	(0.101)
$(\Delta MC \mid \Delta MC>0)_{-1}$	0.103	0.103
	(0.116)	(0.116)
$(\Delta MC \mid \Delta MC<0)$	-0.560***	-0.559***
	(0.0958)	(0.0959)
$(\Delta MC \mid \Delta MC<0)_{-1}$	0.210*	0.211*
	(0.125)	(0.127)
ΔTAX	-0.263***	0.0945
	(0.0634)	(0.0987)
ΔTAX_{-1}	-0.154	-0.173
	(0.0950)	(0.123)
$(\Delta RP \mid \Delta RP>0)_{-1}$	0.0950	-0.256***
	(0.0985)	(0.0633)
$(\Delta RP \mid \Delta RP<0)_{-1}$	-0.167	-0.154
	(0.123)	(0.0952)
Adj R^2	0.9104	0.9101
N of obs	200	200

Variables starred *, ** and *** indicate significance at the 10% 5% and 1% level, respectively. Standard errors in parenthesis.

The results for the regression of equation (3) are displayed in Table 3. The first column shows the OLS regression with *MONTHVOL* among the independent variables. In the second column, where the 2SLS regression is displayed, *MONTHVOL* is replaced by expected monthly volumes obtained from the regression in Table 2. The signs on the estimated coefficients of *RP*, *MC* and *TAX* are the same as in Borenstein and Shepard and the magnitude of the estimates are almost the same, but slightly higher in Sweden. *TIME* is not included in Borenstein and Shepard's regressions, since they have a panel with a dummy for each period.

The estimate of ΔMC_{t-1} is significantly negative for both positive and negative changes in *MC* and the estimate of ΔMC_{t-1} is significantly positive for negative changes in *MC*. Asplund, Eriksson and Friberg (1997) find that part of the price adjustment occurs the period after the change in *MC*. A cost increase will lead to a falling margin in the current month, since only part of the adjustment takes place immediately. In the following month, the margin will increase as some of the price adjustment takes place in this month. This explains the positive sign on ΔMC_{t-1}. The pattern of the response to marginal cost changes is the same in the United States. An augmented Dickey – Fuller test strongly rejects the null hypothesis of a unit root in the residuals.

The variable of interest for testing Haltiwanger and Harrington's model is E(*MONTHVOL*) The punishment of a price war after deviating from an implicit collusive agreement is harder when E(*MONTHVOL*) is high. A hard punishment facilitates implicit collusion, i. e. a high price. The theory would be supported by a significant positive estimate of the effect of E(*MONTHVOL*) on *MARGIN*, but as seen in Table 3, this is not the case.

Part II

The regression in the previous subsection was made to be as comparable as possible to Borenstein and Shepard's investigation of the American market. One possible explanation for the lack of support for the theory might be that the model is not appropriate for the Swedish market. Asplund, Eriksson and Friberg (1999) have investigated the Swedish retail gasoline market. In the next regressions we follow the specification chosen in that paper, but extend the short-run adjustment regression with E(*ΔMONTHVOL*), as estimated in Table 2, among the independent variables in order to capture the possible effects on the price of expected demand changes. Haltiwanger and Harrington's theory is supported if the expected change in volumes has a positive effect on the price.

The Engle – Granger two-step estimation technique is utilized. In the long-run regression *RP* is estimated as a function of *TAX*, *MC* and *WAGE*. *WAGE* is an index of nominal hourly wages in the manufacturing sector in Sweden and is included in order to capture cost changes apart from changes in *TAX* and *MC*.

Table 4: Long-run relationship: Dependent variable *RP*

VARIABLE	
CONSTANT	24.0***
	(6.25)
MC	0.888***
	(0.0307)
TAX	0.662***
	(0.0354)
WAGE	2.00***
	(0.128)
Adj R^2	0.9854
D-W	0.71
N of obs	204

Variables starred *, ** and *** indicate significance at the 10% 5% and 1% level, respectively. Standard errors in parenthesis.

In the short-run regression, *ΔRP* is estimated as a function of cost changes, deviations from the long-run equilibrium, $\hat{u}$, and E(*ΔMONTHVOL*), where the last variable is the one of prime interest. The results are given in Table 5.

Table 5: Short-run relationship: Dependent variable ΔRP

VARIABLE	
CONSTANT	0.823 (0.658)
$\hat{u}$	-0.316*** (0.0625)
ΔTAX	0.770*** (0.116)
ΔTAX_{-1}	-0.0770* (0.0407)
$\Delta WAGE$	0.545 (0.478)
ΔMC	0.550*** (0.0705)
ΔMC_{-1}	0.110* (0.0567)
ΔMC_{-2}	0.0585 (0.0486)
E$(\Delta MONTHVOL)$	-0.00511 (0.0135)
$Adj.R^2$	0.633
D-W	2.05
N of obs	199

Heteroscedasticity consistent standard errors. Variables starred *, ** and *** indicate significance at the 10% 5% and 1% level, respectively. Standard errors in parenthesis.

The results are very similar to those in Asplund, Eriksson and Friberg. An increase in *MC* leads to an increase in *RP*, partly in the current period, partly with a lag and the response to changes in TAX is immediate. The negative sign on $\hat{u}$ implies that RP is adjusted towards the long-run equilibrium as estimated in Table 4. An augmented Dickey – Fuller test strongly rejects the null hypothesis of a unit root in the residuals. The model of Haltiwanger and Harrington is not supported for this specification either, since it predicts a significant positive sign on E$(\Delta MONTHVOL)$. Certain other, here unreported, specifications have been estimated. One of these allows for asymmetric responses to positive and negative changes in *MC*. A regression on real variables has also been estimated. None of the alternative specifications supports the model.

5 Some evidence concerning quantities and margins before and after tax increases

Tax increases induce demand fluctuations since consumers buy more gasoline before and less immediately after a tax increase. These changes in demand in connection with tax increases give another opportunity to study price responses to changes in demand. Table 6 displays deviations from expected sales for periods around tax increases. Expected sales are obtained from a regression with season and GDP as independent variables (the same regression as in Table 2 but without $TAXINC_{t-1}$, $TAXINC$ and $TAXINC_{t+1}$ among the independent variables). Tax changes always occur on the first day of the month. Period t is the first month after a tax increase. The month after a tax increase, sold quantities fall on average 44000 m^3, or about 10%, below what would be expected if there were no tax increase. After the largest tax increases, the fall is about 25%. In the preceding period (i.e. t-1) and the second period following a tax increase (i.e. $t+1$) sales are, on average, increased. There is no clear pattern for more distant periods. The increase in sold volumes in the period preceding the tax increase is as expected, since consumers buy more gasoline before the tax increase. The increase in the second period after the tax increase may be explained by the inventory technology (e. g. it might take somewhat more than a month for many consumers before they must buy gasoline for the first time after a tax increase).

Table 6: Deviations from expected volume of sales 1000 m^3

Date	*Taxchange*	*t-2*	*t-1*	*t*	*t+1*	*t+2*
930115	112,4	-6	76	-133	52	16
841215	50	37	53	-143	94	-33
900115	38	15	5	-60	70	-96
801015	25,1	1	62	-113	1	38
880415	25	2	71	-118	72	-6
870715	24	-60	77	-36	-29	-11
960115	15	-2	-16	22	31	-58
960915	11	30	0	-18	28	-17
820415	7,1	-1	34	-33	-10	42
840115	6	27	-7	-7	24	
840515	6		34	-21	8	26
910715	4	3	-55	42	-7	-5
940115	3	22	-2	-17	0	37
860115	2	-39	10	18	-8	-36
Average		2	24	-44	23	-8

The price is always increased by the same amount as the tax increase at the date the tax change occurs, so there is no change in margins the day the tax increase occurs. Hence, any change in margin in response to changes in demand surrounding the tax increase must come through price adjustments before and after the price change. The margin is measured the fifteenth day of each month. The change in margins as displayed in Table 7 is the difference in margins between the fifteenth day of the present month and fifteenth day of the preceding month.

Table 7: Changes in *MARGIN*

Date	*Taxchange*	*t-2*	*t-1*	*t*	*t+1*	*t+2*
930115	112,4	-4,9	6,0	-10,9	-12,5	-3,5
841215	50	-3,0	-4,6	-7,4	-8,5	5,0
900115	38	1,2	-1,6	10,2	2,4	1,2
801015	25,1	14,1	-0,5	-9,4	-9,4	7,9
880415	25	-5,6	1,7	-26,5	8,3	5,8
870715	24	-0,9	-0,6	-9,1	1,0	23,8
960115	15	5,5	-7,9	-0,8	2,1	-3,3
960915	11	5,2	-3,6	0,2	-10,4	4,3
820415	7,1	1,5	20,4	-19,2	-20,7	0,4
840115	6	-4,4	6,7	4,3	-39,8	
840515	6		-2,0	-1,3	-4,7	1,9
910715	4	7,2	-4,3	-8,5	1,1	-0,1
940115	3	24,6	0,4	-7,6	4,0	6,3
860115	2	15,2	24,5	2,8	16,3	0,1
Average		4,3	2,4	-5,9	-5,0	3,8

One problem with the interpretation of changes in margins is that substantial tax increases may change the optimal margin if, for example, demand elasticity changes. This is probably a minor problem for smaller tax changes. Another problem is that the number of observations is small.

This said, the responses of margins can be compared with the predictions from theory. In period *t-1* demand will be high during the remaining fifteen days of the month. Demand falls in the month of a tax increase. The prediction of the theory is that it would be tempting to cut prices and get a short-run profit since, the short-run profit is higher than usual in the current period and punishment is less severe as demand is low during the punishment phase. According to the theory, the margins must fall in *t-1*, thereby decreasing the profits from deviating, for collusion to remain sustainable. Demand is unusually low in period *t* and expected to increase, so the theory predicts an increase in the margin by a reasoning analogous to the period *t-1*. As seen in Table 7, these predictions are not supported by data, if anything, the opposite seems to be the case. A binomial test for the sign of changes in margins for period *t-1* is insignificant. For period *t* is the opposite of what is predicted by the theory, a decrease in margins, significant at the 10% level.

6 Conclusions

From the regressions in Table 3 and Table 5, we found no support for Haltiwanger and Harrington's model. Expected increases in demand did not have a significant positive effect on current margins, nor did the changes in demand in periods close to tax increases follow the predicted pattern. The question is now: can this lack of support for theory in the Swedish market be explained by some difference between the Swedish and the American gasoline market?

The Swedish and the American market are alike in many respects. Demand follows a similar pattern and the features of the product are, in most respects the same. (E. g. physically relatively undifferentiated between brands and low inventory capacity among consumers.) However, there are some differences that may explain the different price pattern over the seasonal cycle. In Sweden, 95% of the gasoline is sold by the seven largest firms. In the United States, the market is much less concentrated with a large fraction of the sales supplied by independently operated stations. The firms in Sweden try to induce consumer brand loyalty with advertisements and rebates. If consumers have a higher brand loyalty in Sweden, which is reasonable to assume but hard to prove, the short-run gain in sales after a price decrease will be lower than in the United States. Another explanation may be that the market concentration is lower in the United States, at least for consumers that can choose between different sub-markets in different cities, for example long distance travelers. In Sweden, it is much more common that the same competitors are present in all sub-markets, which reduces the number of competitors for consumers present in several sub-markets. This gives rise to two effects which reduces the short-run gain for a firm from reducing its price. There will be fewer competitors for customers who are present in many sub-markets, which reduces the number of customers one firm can get by reducing its price. It also decreases the costs of monitoring the prices of the competitors. It is reasonable to assume that a firm will know the price of those competitors located closely to that firm. This is always the case in the Swedish market, where all major firms compete with each other in many, although of course not in all, sub-markets. In the United States firms may have high costs for monitoring the price of competitors in other sub-markets. Hence, it is reasonable to assume that the deviation period, when the deviating firm makes a higher profit, is shorter in Sweden than in the United States.

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Chapter 3

Essay II: Price Adjustments by a Gasoline Retail Chain

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Price Adjustments by a Gasoline Retail Chain

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Abstract

We use daily data to examine price responses in the Swedish gasoline market to changes in the Rotterdam spot price, exchange rates and taxes. The distribution of price adjustments by a leading retail chain, for the period January 1980 to December 1996, is symmetric with no small adjustments. An error correction model shows that, in the short run, prices gradually move towards the long-run equilibrium in response to cost shocks. There is some evidence that, also in the short run, prices are stickier downwards than upwards. Prices respond more rapidly to exchange rate movements than to the spot market price. Our analysis emphasizes that to fully understand price adjustments it is necessary to examine data sets where the sample frequency at least matches that of price adjustments.

Keywords: Price adjustment; sticky prices; pass-through of costs; gasoline market

JEL classification: *C*22; *E*31; *F*14; *L*71

I. Introduction

In this paper we study the pass-through of cost changes to retail prices in the Swedish gasoline market. This market offers an unusually clean case for testing hypotheses on firms' reactions to changes in the underlying costs of inputs, as both price and costs are observable on a daily basis. Despite daily fluctuations in the spot market price and the exchange rate, the retail price is held constant in the short run. In the longer run, however, prices follow the cost movements quite closely. The first part of the paper provides a detailed description of individual price adjustments. We then proceed to an econometric analysis of the long-run pass-through of costs, in particular the presence of various asymmetries in the pricing pattern.

Figure 1 shows the relationship between price and costs in the Swedish

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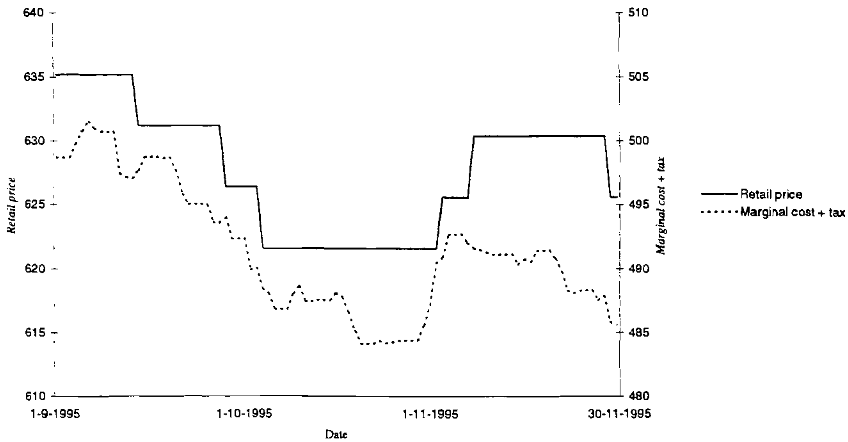

Fig. 1. Development of retail price and marginal cost + tax (September 1995 to November 1995)

gasoline market for September–November 1995. The pattern can be regarded as typical for many products; the price remains fixed for some period of time, and when adjusted it is in the direction motivated by the underlying cost. The stickiness of prices is a well-known phenomenon and has been the focus of considerable theoretical and empirical attention; see e.g. Blanchard and Fisher (1989). Let us use Figure 1 to illustrate some of the problems encountered in empirical studies.

Our work is related to a number of recent empirical studies concerned with price adjustments in response to cost changes. Awh and Primeaux (1992) use annual price data from electric utilities and measure costs by total operating expenditure. Dahlby (1992) uses quarterly data on insurance premiums and Kraft (1995) annual price indices for two-digit German industries, and both control for cost with wage indices. A common feature of these studies is that the data have low frequency, in Figure 1 corresponding to a situation where prices and cost are only observed at long intervals. Moreover, the prices are often average market prices whose development will be smoother than prices set by individual firms.

Another line of research uses quoted prices of specific products and is primarily interested in how often prices are adjusted in times of inflation. Cecchetti (1986) studies cover prices of magazines and Kashyap (1995) prices in retail catalogues. Lach and Tsiddon (1996) provide evidence of the timing of price adjustment of food products and beverages in Israel. With reference to the data depicted in Figure 1, the price is measured accurately but the development of costs is unobserved insofar as it differs from inflation.

Since the data on input and output prices are relatively clean, gasoline prices have been the subject of several studies of price adjustments. Some recent examples include Bacon (1991), Borenstein *et al.* (1997), Borenstein and Shepard (1996a, 1996b), and Slade (1992); additional references are found in Duffy-Deno (1996). However, all studies that examine long time periods use price data aggregated to regional or national levels. Again referring to Figure 1, information is lost concerning when, and by how much, individual firms adjust their prices.

A recurring theme in the literature is whether price adjustments are symmetric with respect to some underlying variable(s). For example, Borenstein *et al.* (1997) are interested in whether gasoline prices are more flexible upwards than downwards. Feenstra (1989) tests whether the long-run pass-through of tariffs and exchange rates onto prices of Japanese cars, trucks and motorcycles in the US market is equal. In this paper we test for various symmetries, e.g. upward and downward flexibility of prices, symmetry in response to exchange rates, spot market prices and taxes.

II. Market and Data Description

We study the retail price of leaded premium gasoline in Sweden for the period 1 January 1980 to 31 December 1996. Gasoline is sold almost exclusively by branded stations and prices in the retail market are determined by the list prices of retail chains (hereafter referred to as "firms"). In this respect the Swedish gasoline market is distinctly different from the market in many other countries, for instance the United States where vertical integration is much less prevalent; see Borenstein *et al.* (1997). The Swedish gasoline market was remarkably stable during the sample period. For example, the quantity of gasoline sold annually increased only slightly over the period (from 4.9 million m^3 in 1980 to 5.7 million m^3 in 1996) and market concentration was virtually unchanged (the Herfindahl index was 0.127 in 1980 and 0.152 in 1996). This is attractive since changes in these two factors could influence pricing behaviour; see Borenstein and Shepard (1996b) for discussion and evidence.

From January 1990 and onwards the data set includes prices for virtually all gasoline retail chains (seven firms representing 94.5 percent of sales in 1996). For the period 1980–1989 price data are available for two of the firms (Shell and Norsk Hydro). We restrict our attention to the retail list price of one of the firms, Shell, which throughout the period was either the largest or the second largest firm (with a market share of 16.5–21.0 percent; source: Swedish Petroleum Institute). The behaviour of Shell's retail price is representative of the prices of other firms since, more often than not, firms' list prices are identical and they all adjust their prices either the very same day or within a day.[1] This can be exemplified by two simple observations. During the 17-year period, there were just 11 occasions when the prices of Shell and Norsk Hydro were different more than three days in a row. And in the seven-year period 1990–1996, the saving from buying 1500 litres/year from the firm with the lowest average price, compared to the firm with the highest, was a meagre SEK140 (or less than US$ 20!).[2] As we are interested in how the price is adjusted in response to cost changes, and not strategic issues such as price leadership, the fact that prices of firms sometimes differ one or two days is of no importance to our analysis.[3]

Even though the prices are more or less identical for a retail chain across the country (except for some minor differences due to transport costs) there is some local variation in retail prices. The local variation refers primarily to constant differences in price levels (for instance, due to localization of the stations) and not to differences in the pattern of price adjustments. The chain from input price to consumer price is simple in Sweden. The relevant input price is the Rotterdam spot price for gasoline. Some firms buy their gasoline at this price on the spot market, but even those who operate their own refineries claim to use it as the transfer price between the producing and the selling divisions.

Let RP denote the retail list price of one litre of premium leaded gasoline (excluding the value-added tax introduced in Sweden in March 1990) measured in SEK$*$100; see the Appendix for detailed descriptions of the variables. This price is observed every day and may be adjusted any day. The Rotterdam spot market price of gasoline in US$, SP, and the SEK/US$ exchange rate, E, are not quoted on weekends and public holidays. We assume these variables to be unchanged on these days. The marginal cost of gasoline in SEK is denoted MC and defined as $SP*E$. The only variable not measured in SEK (or US$) per litre is $WAGE$, an index of the nominal wage. The price changes are denoted ΔRP. The change in marginal cost since the last price adjustment is denoted ΔMC. With marginal cost composed of two

[1] According to discussions with Shell, the procedure for adjusting the price is as follows. Executives meet before noon and decide whether or not to adjust and, if so, the magnitude. The suggested retail price is then submitted by fax to all the gasoline stations (612 as of January 1996) which adjust the price the following morning. It is not possible to keep a price adjustment secret from rival firms until the next morning. News agencies are also notified and transmit the information to rivals. According to Shell, the cost of conveying the message has decreased over time as a consequence of new information technology.

[2] The saving would be larger, SEK 2,817 $\sim$ US$ 350 (or about 3.5 percent), if the discount gasoline retailer JET had been included. JET, however, is only active in some regions and has unmanned stations.

[3] The complexity involved in econometric modelling of a dynamic game along the lines of Slade (1998) is quite substantial and would cloud the present issue of price responses to cost changes.

factors, *SP* and *E*, it is relevant to decompose ΔMC into its parts. $\Delta E{*}SP$ denotes movements due to the exchange rate, holding *SP* fixed. $\Delta SP{*}E$ is defined analogously. A factor of great, but varying, importance is the quantity (per litre) tax, denoted *TAX*. In 1980, 1982, 1988 and 1996 it accounted for approximately 50, 40, 65 and 50 percent of *RP*, respectively. A legally imposed price freeze and three price wars are excluded. For the purpose of this work it is obvious that those periods when price changes do not reflect cost movements should be excluded.[4]

III. Price Adjustment: Part 1

Much of the recent literature on pricing has been concerned with the distribution and frequency of price adjustments. To add to the evidence on the behaviour of prices, we begin with a detailed description of the distribution of price adjustments, before relating them to cost changes. There are 250 price adjustments in the sample, implying that on average price was adjusted every third week. The size distribution of price increases is strikingly similar to that of price decreases, as shown in Table 1a and Figure 2. A stark feature of the figure is the minimum absolute size of $|\Delta RP| = 2$. This strongly indicates that there is some fixed cost associated with price changes that keeps firms from making very small adjustments. All adjustments larger than $|18|$ are associated with tax changes or large discrete changes in *E* due to devaluations of the Swedish currency.[5] The next important fact is the existence of absolute changes of quite different magnitude in the interval between $|2|$ and $|18|$, with changes around $|6|$ being the most frequent.

Considering that the data cover 17 years, many with high general inflation,

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[4]The price freeze period excludes 29 August 1986 to 23 August 1987. The price war periods exclude 2 January 1984 to 21 February 1984, 2 April 1988 to 6 May 1988 and 30 September 1993 to 28 October 1993 and were identified partly from information provided by firms. We dropped from our sample the period beginning the day after the last price adjustment prior to the price war (freeze) and ending the day after the price increase that ends the price war (freeze). Since the regressions include at most two lagged variables, this implies that at most two additional observations were dropped for each occasion. The coefficients in the reported regressions are, however, only marginally affected by excluding the price war periods. Ellison (1994) and Slade (1992) have dealt with the causes and timing of price wars. Inspection of the three episodes in our sample revealed no apparent explanation (such as unusually high or low margins) as to why they occurred.

[5]Until May 1991 the US$ was included in the currency basket against which the SEK was tied. Even though the US$ was included in the basket, the SEK/US$ exchange rate exhibited large fluctuations. From May 1991 to November 1992, the SEK was tied to the ECU. Since then it has floated.

Table 1a. *Descriptive statistics for all price adjustments*

	Mean	St. Dev.	Skew	Kurt	Min.	Max.	Median	Cases
$\Delta RP \neq 0$								
ΔRP	1.55	11.05	3.0	23.7	−16.20	95.20	3.20	250
ΔMC	0.04	9.26	0.3	6.0	−37.26	35.54	−0.27	250
$(\Delta SP{*}E)_{-1}$	−0.31	8.84	0.4	6.2	−35.09	35.42	0.00	250
$(\Delta E{*}SP)_{-1}$	0.37	3.94	1.4	15.2	−14.27	26.24	0.12	250
DAYSFIXED	21.59	23.90	2.6	12.4	0	177	13	250
$\Delta RP > 0$								
ΔRP	9.09	10.22	5.4	41.8	2.00	95.20	6.00	130
ΔMC	4.62	8.69	0.9	5.3	−17.00	35.54	3.98	130
$(\Delta SP{*}E)_{-1}$	3.11	8.37	0.9	5.5	−17.64	35.42	1.79	130
$(\Delta E{*}SP)_{-1}$	1.54	4.46	1.4	14.0	−14.27	29.24	0.70	130
DAYSFIXED	24.81	26.64	2.6	12.3	1	177	14	130
$\Delta RP < 0$								
ΔRP	−6.62	3.59	−1.4	3.9	−16.20	−2.00	−5.60	120
ΔMC	−4.92	7.05	−1.2	8.5	−37.26	21.14	−3.55	120
$(\Delta SP{*}E)_{-1}$	−4.02	7.38	−0.8	7.1	−35.09	24.02	−2.78	120
$(\Delta E{*}SP)_{-1}$	−0.89	2.65	−0.5	4.4	−9.77	5.33	−0.68	120
DAYSFIXED	18.10	20.06	2.2	7.8	0	104	11	120

Table 1b. *Descriptive statistics for the subperiods 1980–1984, 1985–1989 and 1990–1996*

	$\lvert\Delta RP\rvert$ Mean (st. dev)	$\lvert\Delta MC\rvert$ Mean (st. dev)	$\lvert\Delta SP*E\rvert$ Mean (st. dev)	$\lvert\Delta E*SP\rvert$ Mean (st. dev)	*DAYSFIXED* Mean (st. dev)	$\Delta RP > 0$ Nobs.	$\Delta RP < 0$ Nobs.
1980–1984	8.89 (8.54)	7.61 (7.33)	7.10 (6.61)	4.08 (4.79)	38.18 (27.49)	28	16
1985–1989	9.45 (5.13)	9.06 (7.57)	8.63 (7.99)	3.30 (3.30)	31.84 (36.30)	20	24
1990–1996	7.22 (8.23)	5.53 (5.81)	4.93 (5.81)	1.88 (2.05)	14.30 (13.13)	82	80
1980–1996	7.90 (7.86)	6.52 (6.56)	5.96 (6.52)	2.52 (3.06)	21.59 (23.90)	130	120

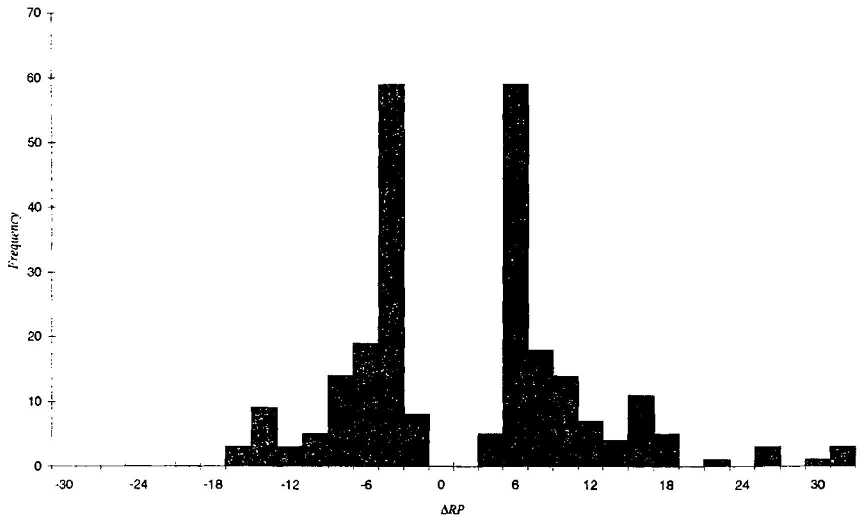

Fig. 2. Frequency distribution of price adjustments (rightmost class contains all price adjustments greater than 30)

one would expect a distribution skewed to the right, not the symmetric distribution seen in the figure. This has two explanations. First, the retail price of gasoline is largely determined by the highly volatile spot market price and the likewise volatile exchange rate, which incidentally resulted in the fact that the input cost in local currency was almost the same in 1980 as in 1996. Second, the size of the price adjustments and the time between them are not stable over time. Table 1b shows that there are more adjustments, which on average are smaller in size, in the 1990s than in the 1980s. Nevertheless, there is considerable variability in the size of adjustment within each subperiod.

There is some seasonal variation in sales with a peak in the summer months, as seen in Table 2. If there are fixed costs of adjusting price, but the loss from a misadjusted price is proportional to the sales volume, one would expect to find more adjustments during the high demand months. This is not borne out by the table. On the other hand, we find substantial variation in the frequency of price adjustments over the week; few changes made on Saturdays, Sundays and Mondays. The simplest explanation is that financial markets and the Rotterdam spot market are closed on Saturdays and Sundays and thus no new information arrives on these days.

Now turn to an examination of the relationship between price and cost. As a first step we use monthly data to estimate the long-run relationship in

Table 2. *Number of price adjustments, by month and day of the week*

	Jan.	Feb.	Mar.	Apr.	May	June	July	Aug.	Sep.	Oct.	Nov.	Dec.
Quantity 1000 m^3	370	386	438	446	465	497	526	501	450	448	422	433
$\Delta RP > 0$	12	13	10	17	5	13	9	12	11	10	7	11
$\Delta RP < 0$	9	6	9	6	7	11	10	10	13	14	11	14
	Mon.	Tue.	Wed.	Thu.	Fri.	Sat.	Sun.					
$\Delta RP > 0$	9	31	33	29	18	5	5					
$\Delta RP < 0$	13	34	25	20	20	8	0					

levels between *RP* and the explanatory variables *MC*, *TAX* and *WAGE*. Figure 3 shows the development of the first three over the period.[6] The margin (*RP–MC–TAX*) is increasing over the period. This reflects general inflation, which is highly correlated with the explanatory variable *WAGE*. This regression fills a dual purpose. First, a high explanatory value of the regression indicates that these three variables are able to explain almost the entire variation in price. Second, the estimated error term was used in error correction regressions, reported below. The estimated relationship is

$$\underset{}{RP} = \underset{(6.26)}{22.9} + \underset{(0.03)}{0.902} * MC + \underset{(0.03)}{0.658} * TAX + \underset{(0.12)}{2.01} * WAGE + u. \quad (1)$$

$$\text{DW: } 0.63 \qquad \text{Adj. } R^2\text{: } 0.988$$

In the long run, there is a near one-to-one relationship between *RP* and

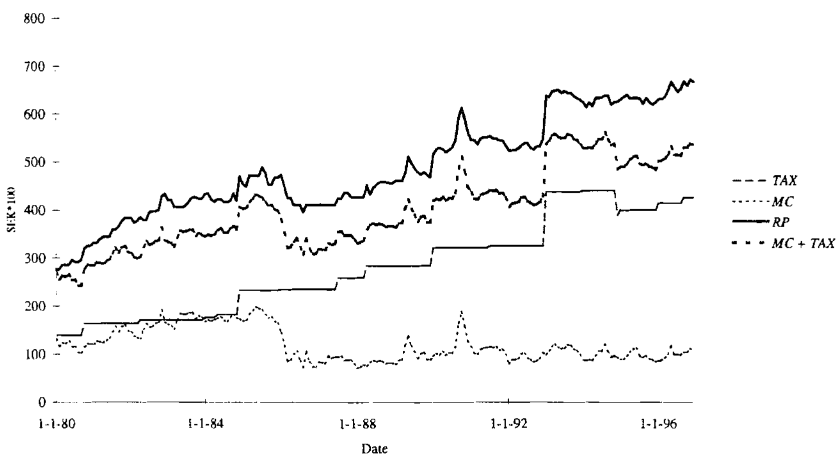

Fig. 3. Development of *RP*, *MC* and *TAX* over the period 1980–1996

[6]A log formulation is sometimes used in studies of pass-through of costs when only price indices are available; see e.g. Feenstra (1989). However, with a quantity tax that constitutes a large and varying share of the retail price, the coefficients from a log formulation are less readily interpreted. Moreover, the specification in levels is standard in the literature that examines price adjustment on gasoline markets; see, for instance, Borenstein *et al.* (1997). We also estimated all reported regressions in log form (available on request). The results are robust and none of the main conclusions are altered.

MC. *TAX* and *WAGE* are highly correlated with each other as well as with inflation, which may go some way towards explaining the low coefficient on the former.[7] The Durbin–Watson statistic shows evidence of positive autocorrelation – periods with relatively high prices tend to be followed by another period of relatively high prices, and vice versa.

A non-negative relation between recent cost changes and price adjustments is expected. Our data suggest that this is true on average, but that there are considerable deviations on many occasions. In Table 1a the means and medians of ΔMC have the expected signs for $\Delta RP > 0$ and $\Delta RP < 0$, respectively. However, the minimum and maximum values of ΔMC show that sometimes the price moves in one direction when the marginal cost has moved in the other direction. In fact, about one-fifth of the price adjustments are in these two quadrants, but in most of these cases the preceding change in marginal cost has been small. Although this could be due to omitted variables in the marginal cost measure this is probably not the case here, since we would then arguably get far lower explanatory power in regression (1).[8] It could also be explained by strategic elements. For example, we have not identified all price war periods, the firm may wish to undercut its competitors, or it may react to its competitors' previous price changes. We have looked thoroughly for more price wars but found none. The latter two explanations can only be examined for the 1990s where we have price information from all firms. The observations in the "wrong" quadrants could not be related to instances when Shell had a price different from its competitors. Furthermore, we could get the observed pattern if the firm tried to predict marginal cost movements and some price changes were made in anticipation of future developments.[9] Hedging opportunities and inventory situation could also influence the price-setting process. We explicitly asked the firm about these issues and it clearly stated that such concerns were not present in the pricing decision.

An explanation for the large number of observations of the "wrong" sign

[7]Excluding *WAGE*, the estimated coefficient on *TAX* is above unity as it captures not only the quantity tax but also inflation over the period.

[8]As an anonymous referee suggested, transport costs are excluded and may influence the results. Oil-tanker freight rates show wide fluctuations over the years. However, they are at least partly incorporated in our measure of marginal cost (FOB Rotterdam). The change in road transport cost from terminal to station should be positively correlated with our measure of the change in marginal cost.

[9]Asset prices in liquid markets (such as the foreign exchange market and spot market for gasoline) follow random walks, and are thus intrinsically unpredictable. Dickey–Fuller tests for random walks in *E*, *SP* and *MC* (203 monthly observations) confirm this presumption. The test statistics are (including a constant) −2.24, −2.87 and −2.22, respectively, compared to a critical value of −3.51.

is a gradual adjustment pattern. If the firm, for some reason, wishes to avoid large price changes, this may result in occasional violations of a monotone relation between cost change and price adjustment. Take the following simple example: marginal cost rises sharply by 15, but the firm judges that this is too large a price increase to make at once. It instead raises the price by 7 in a first step, such that the price is still 8 "too low". Marginal cost now falls by 1. The firm again raises its price by 7, even though cost has actually fallen since the last price adjustment.

To capture the possibility that price adjustments partly reflect previous changes in costs, lagged variables were included in the subsequent econometric analysis. We estimate the relationship between all the price changes and the preceding changes in marginal cost and tax. *WAGE* was not included since it is only observed monthly and price adjustments do not have a fixed frequency. In addition, we split the observations into positive and negative price adjustments.[10] The results are reported in Table 3.

When pooling all observations (column 3:1) the intercept is not significantly different from zero and taxes are reflected almost completely in the price change. About half of the drift in marginal cost since the last price adjustment is reflected in today's price change. The pricing decision also reflects earlier ΔMC with a coefficient of 0.28. Adding the two coefficients yields approximately the long-run coefficient in (1) – it takes two adjustments to respond to the cost movement. The picture is distinctly different in 3:2 and 3:3 where we condition on the sign of the price change. In absolute terms, the size of the intercept is the same, but in contrast to 3:1 significantly different from zero. This mirrors the minimum absolute size of price changes. Hence, pooling all price changes results in a bias towards a steeper slope of the regression line compared to the slopes conditional on the direction of the adjustment. The coefficient on ΔMC is greater (0.28) for price rises than for price cuts (0.19); indicating an asymmetry in the price adjustments. The effect of lagged ΔMC remains significant but is considerably lower than in 3:1. It is noteworthy that price increases are well explained by the changes in the underlying costs, compared to the price cuts. This is

[10]The days when price is adjusted presumably do not represent a random sample of days. Moreover, days when the price is increased do not reflect a random draw from the days when price was adjusted. Both factors can potentially cause sample selection bias. To test this we used a method similar to the well-known Heckman two-stage procedure. The first stage selection mechanism is an ordered (rather than a standard binary) probit model and the second stage, the price change regression, is an ordinary least squares estimator with a correction for sample selection. However, the price is held constant on about 95 percent of the days and the predictive ability of the sample selection model is poor. Reducing the frequency to weekly data improves predictive power somewhat, but insufficiently to allow identification of the effects of sample selection. For a full account of the results, see Asplund *et al.* (1997).

Table 3. *Price adjustments in relation to changes in marginal cost, spot market price, exchange rate, and tax; individual price adjustments (no fixed frequency)*

Variable	3:1 ΔRP	3:2 $\Delta RP \vert \Delta RP > 0$	3:3 $\Delta RP \vert \Delta RP < 0$	3:4 ΔRP	3:5 $\Delta RP \vert \Delta RP > 0$	3:6 $\Delta RP \vert \Delta RP < 0$
CONSTANT	0.0405	5.51**	−5.40**	0.295	5.37**	−5.47**
	(0.366)	(0.415)	(0.380)	(0.363)	(0.420)	(0.395)
ΔTAX	0.887**	0.801**		0.885**	0.802**	
	(0.0426)	(0.0298)		(0.0419)	(0.0297)	
ΔMC	0.548**	0.283**	0.192**			
	(0.0392)	(0.0403)	(0.0430)			
$\Delta SP*E$				0.513**	0.277**	0.193**
				(0.0406)	(0.0417)	(0.0428)
$\Delta E*SP$				0.790**	0.377**	0.134
				(0.0909)	(0.0793)	(0.121)
ΔMC_{-1}	0.279**	0.169**	0.123**			
	(0.0403)	(0.0398)	(0.0346)			
$(\Delta SP*E)_{-1}$				0.279**	0.165**	0.120**
				(0.0418)	(0.0424)	(0.0363)
$(\Delta E*SP)_{-1}$				0.260**	0.177*	0.131
				(0.0923)	(0.0845)	(0.0847)
Adj R^2	0.733	0.853	0.178	0.742	0.855	0.171
DW	2.21	1.64	1.21	2.18	1.59	1.21
Nobs	249	129	120	249	129	120

Notes: Heteroscedascity-consistent standard errors. ** denotes variables significant at the 1 percent level and * at the 5 percent level.

partly due to tax changes, but even excluding those observations the explanatory power is greater for price increases.

Columns 3:4–3:6 show the results from the decomposition of marginal cost into movements in the spot price ($\Delta SP{*}E$) and the exchange rate ($\Delta E{*}SP$). The coefficients for the constant and the tax change are virtually the same as in 3:1–3:3. Each of the other coefficients has the expected sign. The interesting point, however, is that the price response to a given change in the marginal cost is dependent on whether it is caused by the exchange rate or the spot market price. Pooling all observations (3:4) reveals that exchange rates are more important for the adjustment than spot market prices (the current coefficient is considerably larger, whereas the lagged is about the same, and an F-test confirms that their sums are statistically different). Closer examination of columns 3:5 and 3:6 shows that the strength of the exchange rate pass-through comes primarily from the price increases, whereas changes in spot market prices are important for both rises and cuts.

IV. Price Adjustment: Part 2

Even though lagged independent variables were included in the analysis in the preceding section, the standard procedure for studying dynamic adjustment is to use an error correction framework, which is pursued here. Instead of using the change in the cost variables since the last price adjustment, we now analyse cost and price changes on a monthly basis. The first step in an error correction approach is to estimate the relationship in levels between price and the explanatory variables, reported in equation (1). The estimated error term, $\hat{\boldsymbol{u}}$ may be interpreted as the deviation from the long-run relation. Engle–Granger tests reject that $\hat{\boldsymbol{u}}$ has a unit root (the test statistic is, including a constant, -6.60 compared to a Dickey–Fuller critical value of -3.51) such that (1) represents a cointegrating relationship. The second step is estimation of the error correction form

$$\Delta RP_t = \alpha_0 + \alpha_1 \hat{\boldsymbol{u}}_{t-1} + \alpha_2 \Delta \mathbf{X}_t + \alpha_3 \Delta \mathbf{X}_{t-1} + \alpha_4 \Delta \mathbf{X}_{t-2} + v_t, \qquad (2)$$

where preliminary regressions revealed that two lags of the independent variables are sufficient to capture the dynamics. Note that the Δ's now measure change on a monthly basis. α_1 is an estimate of the fraction of the misadjustment (relative to long-run equilibrium) in the previous month that is "corrected" in the current month.

Overall in Table 4, contemporaneous explanatory variables are significant at the 1 percent level whereas lagged variables have varying significance levels. First, we focus on the estimation reported in column (4:1). A change

Table 4. *Price changes in relation to changes in marginal cost, spot market price, exchange rate and tax; error correction model; monthly frequency*

Variable	4:1 ΔRP	4:2[a] ΔRP	Variable	4:3 ΔRP	4:4[a] ΔRP
CONSTANT	0.717 (0.599)	0.736 (1.01)	*CONSTANT*	0.408 (0.589)	0.102 (1.346)
$\hat{u}$	−0.275** (0.0523)	−0.274** (0.0532)	K	−0.292** (0.0520)	−0.288** (0.0538)
ΔTAX	0.745** (0.0515)	0.750** (0.0512)	ΔTAX	0.726** (0.0520)	0.723** (0.0523)
ΔTAX_{-1}	−0.0591 (0.0525)	−0.0586 (0.0522)	ΔTAX_{-1}	−0.0951 (0.0525)	−0.0907 (0.0529)
$\Delta WAGE$	1.01* (0.452)	0.939* (0.450)	$\Delta WAGE$	1.01* (0.440)	0.898* (0.447)
ΔMC	0.541** (0.0549)		$\Delta SP{*}E$	0.476** (0.0564)	
			$\Delta E{*}SP$	0.850** (0.123)	
$\Delta MC \mid \Delta MC > 0$		0.695** (0.0970)	$\Delta SP{*}E \mid \Delta SP > 0$		0.624** (0.101)
			$\Delta E{*}SP \mid \Delta E > 0$		0.940** (0.172)
$\Delta MC \mid \Delta MC < 0$		0.349** (0.103)	$\Delta SP{*}E \mid \Delta SP < 0$		0.298** (0.108)
			$\Delta E{*}SP \mid \Delta E < 0$		0.745* (0.305)
ΔMC_{-1}	0.183** (0.0586)		$(\Delta SP{*}E)_{-1}$	0.194** (0.0608)	
			$(\Delta E{*}SP)_{-1}$	0.208 (0.124)	

Table 4. *Continued*

Variable	4:1 ΔRP	4:2[a] ΔRP	Variable	4:3 ΔRP	4:4[a] ΔRP
$(\Delta MC \mid \Delta MC > 0)_{-1}$		0.0289	$(\Delta SP{*}E \mid \Delta SP > 0)_{-1}$		0.0660
		(0.103)			(0.110)
			$(\Delta E{*}SP \mid \Delta E > 0)_{-1}$		0.235
					(0.178)
$(\Delta MC \mid \Delta MC < 0)_{-1}$		0.311**	$(\Delta SP{*}E \mid \Delta SP < 0)_{-1}$		0.282**
		(0.107)			(0.108)
			$(\Delta E{*}SP \mid \Delta E < 0)_{-1}$		−0.0368
					(0.288)
ΔMC_{-2}	−0.0643		$(\Delta SP{*}E)_{-2}$	−0.112*	
	(0.0551)			(0.0566)	
			$(\Delta E{*}SP)_{-2}$	0.166	
				(0.126)	
Adj. R^2	0.693	0.698	Adj. R^2	0.711	0.711
DW	1.96	1.97	DW	1.91	1.93
Nobs	178	178	Nobs	178	178

Notes: Heteroscedascity-consistent standard errors. ** denotes variables significant at the 1 percent level and * at the 5 percent level.
[a]Two lags of the signed variables included but coefficients not reported; none significant at the 10 percent level.

in marginal costs of one unit leads to a price adjustment of 0.54 in that same month (which is similar to the point estimate in 3:1). Not all adjustment takes place in the same month; the price change on average reflects 0.18 of the previous month's change in marginal cost. Furthermore, the estimated coefficient for $\hat{u}$ indicates that about 0.27 of the misadjustment from the previous month is corrected. Estimates confirm the finding in the preceding section that tax changes are immediately passed through to prices (point estimate 0.74). It cannot be rejected that the accumulated pass-throughs of ΔMC and ΔTAX are equal. For more details on joint tests and alternative specifications, see Asplund *et al.* (1997).

Next, in column (4:2), we study whether the responsiveness of price is symmetric in decreases and increases of the marginal cost. A rise in MC of one unit yields a contemporaneous price increase of about 0.7. There is no statistically significant effect of the previous month's increases in MC on the price change. A fall in MC of one unit, on the other hand, results in a contemporaneous price effect of only 0.35. However, there is a statistically significant effect of the previous month's decrease in MC on the price change of about 0.31. Taken together, this means that the price responded more rapidly to cost increases than to decreases, but that the accumulated pass-through is symmetric (around 0.7). At the 5 percent level it cannot be rejected that the accumulated effects are the same.

Despite the marked differences in market structure we find a similar pattern in our Swedish data as previously found in British and US data. Borenstein *et al.* (1997) found that, during 1986–1992, US retail gasoline prices responded more rapidly to cost increases than to cost decreases. According to Bacon (1991), British retail gasoline prices responded more rapidly to upward movements in costs during 1982-1989. Kirchgässner and Kübler (1992), on the other hand, found that for the German gasoline market, retail price adjustment had been rapid, symmetric and full to Rotterdam spot price fluctuations during 1980–1989 (but asymmetric 1972–1979).

Column (4:3) presents the results with MC separated into E and SP. Note first that point estimates of $\hat{u}$, ΔTAX and $\Delta WAGE$ are essentially unchanged. The contemporaneous pass-through is greater for exchange rates than for spot prices. It is possible to reject that the accumulated pass-throughs of $\Delta SP{*}E$ and $\Delta E{*}SP$ are the same. At the 10 percent level of significance it can be rejected that the long-run pass-through of ΔTAX is the same as for $\Delta SP{*}E$ and $\Delta E{*}SP$. For taxes and exchange rates, more of the adjustment takes place in the same month compared to the spot market price.

Finally, column (4:4) separates cost changes into increases and decreases in E and SP, respectively. Increases in SP are passed through in the same month as they occur (point estimate of about 0.62), whereas long-run price adjustment to decreases in SP is about the same, but distributed over two months (pass-through of 0.3 in the same month and of 0.28 from the previous month). Similarly, the adjustment to E takes place fully the same month as it occurs. Also in the case of E, the price response to increases is larger (0.94) than the response to decreases (0.74).

V. Discussion

Seventeen years of daily input and output prices for a major gasoline retail chain provide an unusual opportunity to confront price-setting theories with data. The three main theories for how prices are set when there are costs of adjustment are state-dependent pricing with fixed adjustment costs (Ss-models), time-dependent pricing and partial adjustment. Below we summarize the contribution of each of these theories to our understanding of the observed pattern of price adjustment.

At first glance, the Swedish gasoline industry seems to offer a textbook example of state-dependent pricing in its simplest form. Even though input prices move virtually every day, we observe infrequent price adjustment and no small price changes. The price adjustment to tax changes is also consistent with fixed adjustment costs; adjustment appears to take place immediately and almost fully. Evidently there is a fixed-cost component of price adjustments for the retail chain, at a minimum the cost of disseminating the information to all gas stations. The most obvious inconsistency with fixed costs is that on several occasions, the price is increased (decreased) despite the fact that cost has decreased (increased). Other features left unexplained are extended periods when the price is well above the long-run equilibrium level, and that prices are not fully adjusted to this level when they are reset.

We study the timing of the price changes and find only very limited evidence of time-dependent pricing, that is of price adjustment taking place at a specific point in time ("prices are reset on the first day each month"). Time-dependent price-setting rules are likely to be most relevant when it is costly to learn about the state and when one wants to economize on information collection; see Blanchard and Fischer (1989, p. 413). The most important factors in the gasoline market are readily observable and change rapidly, and the firm continuously monitors the market for significant changes in underlying conditions.

If the costs of adjusting price are convex in the size of the price adjustment, the result will be gradual movement towards long-run equilibrium following a shock. This is well in line with the observed response to changes in input cost. However, the full and rapid pass-through of tax changes is inconsistent with adjustment costs that are convex in the price change *per se*. Rotemberg (1982) motivated convex adjustment costs by consumer dislike for large price changes. Instead, this may be interpreted as an aversion to large price changes due to reasons that consumers do not fully understand, whereas "well motivated" price changes are not perceived as equally bad. Tax increases are given extensive media coverage and cannot be blamed on the firm. This behavioural assumption is consistent with both the gradual pass-through of marginal cost changes and the instantaneous pass-through of taxes.

None of the above theories is consistent with all features of price adjustments. This in itself is to be expected, since there are specificities of each single price adjustment that are unobservable to the researcher. A good example of such unobservables concerns beliefs of the firm regarding the price responses of competitors. In the paper we have not studied any strategic considerations by the firm, but they are clearly important for price setting in the short run. We have discussed the issue with representatives of the firm. They pointed out that although all price adjustments can be observed almost instantaneously, rivals have the option of sticking to their old prices or changing them by a different amount. This was given as motivation for avoiding 'very large' price changes, even though they may be called for due to large changes in the observable underlying variables. These behavioural assumptions would also be consistent with gradual adjustment. Again, for a tax change, there is a natural focal point since such changes are known in advance. As the firm said, "everybody knows that tax changes are passed through fully and at once".

This indicates that the characteristics of a variable, such as its stochastic properties, matter for the short-run price response. Tax increases are known in advance, and there is essentially a zero probability that they will fall shortly after. On the other hand, marginal cost is stochastic and there is always the probability that it will revert, such that there is an option value of postponing a price adjustment. This may also go some way towards explaining the marked difference in the price response to exchange rates and spot market prices. The volatility in marginal cost is to a greater extent due to fluctuations in the spot market price than in the exchange rate. Even though both variables enter the marginal cost symmetrically, firms may wait and see whether the spot market price reverts, but react faster to the less volatile exchange rate.

Finally, referring back to Figure 1, inferences on price-setting behaviour are highly dependent on the data analysed. For example, with access only to the price series (or only the distribution of price adjustments), it would be easy to conclude that an Ss-model is an accurate description of price setting. On the other hand, if data were sampled only at (longer) discrete intervals, the conclusion could instead be in favour of a partial adjustment model. This clearly illustrates the necessity of examining data sets where the frequency at least matches the frequency of price adjustments.

Appendix. Variable Definitions and Data Sources

Variable	Definition and data source
RP	Retail list price of premium leaded gasoline for one of the firms, Shell. The retail price is at 0-zone, i.e., where the transport cost is the lowest. VAT (introduced in 1990) is excluded from the price according to the formula: RP = consumer price/(1 + VAT) measured in SEK*100/litre. Source: Shell.
SP	The spot market price of premium leaded gasoline. From November 1985 to December 1996 it is the average of high (from refinery) and low (stored or blended material) daily quotations, FOB Rotterdam for Premium 0.15 g/l. Size is barges (1000–5000 mt) and delivery is North-west Europe. Prices are assessed daily up until 1830 GMT. For the period January 1980 to October 1985 it is for Mediterranean/Italian delivery of regular gasoline plus US$ 24, which was the average difference between the two prices for the period where we have access to both. The FOB price is converted from US$/mt to US$/litre by the factor 8.35*159. *SP* is measured in US$*100/litre. Source: Platt's, London.
E	The exchange rate between Swedish kronor and US dollars. Source: Findata (January 1983 to December 1996), Stefan Nydahl, Uppsala University (January 1980 to December 1982).
MC	Spot market price of premium leaded gasoline measured in SEK*100/litre. *MC* is obtained by: $MC = SP^{*}E$.
TAX	Total tax per litre premium leaded gasoline. Source: The Swedish Petroleum Institute, Annual Report 1995.
VAT	Value-added tax. Calculated on producer price including *TAX*. A *VAT* of 23.46% was levied on gasoline on 1 March 1990; it was subsequently (1 July 1991) increased to 25.00%. Source: The Swedish Petroleum Institute, Annual Report 1995.
WAGE	Index of nominal hourly wages in the manufacturing sector (SNI 3). Source: International Financial Statistics, heading 14465.zf.
ΔX_{lag}	Changes in the variable X are denoted ΔX. In Section III, Δ refers to the change since the last price adjustment and in Section IV it refers to the change since the previous month. The subscript on X denotes the number of lags.
DAYSFIXED	The number of days *RP* has been fixed since the latest price adjustment.

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Chapter 4

Essay III: Prices, Margins and Liquidity Constraints: Swedish Newspapers 1990-1996

Prices, Margins and Liquidity Constraints:

Swedish Newspapers 1990-1996*

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Abstract

Using firm level data from the Swedish daily newspaper industry, we test theories on the effects of liquidity constraints on prices in markets with consumer switching costs. The newspaper industry is of particular interest, since firms set prices in two markets, the subscription market, where switching costs are high, and the advertising market, where switching costs are low. When Sweden enters a recession, we find a relative increase in subscription prices and margins for liquidity constrained firms, which is not the case for advertising prices.

Key words: Liquidity constraints; switching costs; price adjustment; newspaper industry.
JEL classification: D43; E32; G33; L82.

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1 Introduction

Several hypotheses have been advanced to explain how price-cost margins vary over the business cycle. In particular, economists have provided a number of reasons why firms have the incentive to keep prices relatively low in times of high demand and vice versa, which would tend to amplify economic fluctuations. These explanations can broadly be classified into three categories. First, recessions are often accompanied by a tight credit market and some firms face liquidity constraints. Provided that consumers have switching costs, firms can boost short-run profits by increasing prices to avoid default at the cost of foregone future profits, see e.g. Chevalier and Scharfstein (1996). Second, periods of high demand are associated with an inflow of new, unattached customers. In order to capture these for the future, firms use aggressive pricing behavior, Bils (1989). Third, firms gain more by deviating from fully collusive prices in high demand states than when demand is low. To sustain an implicitly collusive equilibrium, the prices in high demand states must be lower than fully collusive prices, Rotemberg and Saloner (1986). These hypotheses and empirical tests on intra-industry data are discussed in more detail below.

To test the hypotheses, we provide an empirical examination of price adjustments in an industry having experienced times with radically different macroeconomic conditions: the Swedish newspaper industry in 1990 to 1996. In 1990, Sweden entered a deep recession, with a falling real GDP for three consecutive years, a quadrupling of unemployment, and a tightening of credits which leading to a threefold increase in bankruptcies. A strong recovery began in 1994 with three years of high/medium-high GDP growth and a gradual loosening of credit. Such broad changes in macroeconomic conditions may, at a first approximation, affect all newspapers to an equal extent, and the question is therefore if they all responded in a similar manner. We test whether the patterns of price adjustments for subscriptions and advertising space depended on local and/or newspaper specific factors. The local factors for which we control are demand factors, such as the development of income and population. Newspaper specific factors include market position and, in particular, liquidity constraints.

Our most important result is that newspapers with weak financial standings showed the highest increases in subscription price during the recession, which improved their margins in the short run. Financial variables could not explain differences in the price increases for advertising space, however. This suggests that newspapers with liquidity constraints attempted to raise short-term profits by exploiting readers, who are believed to have high switching costs, but could not improve their profitability at the expense of advertisers, who are less attached to a particular newspaper. We find no significant differences between monopoly and competitive markets. Finally, our results suggest that in the recovery period, the development of prices and margins were due to unobservable idiosyncratic factors.

2 Theories and Tests of Pricing over the Business Cycle

A priori price-cost margins are expected to be pro-cyclical: in times of high demand, prices are also high. However, it is conceivable that margins are counter-cyclical, or at least have a counter-cyclical tendency. This possibility has spurred a significant empirical literature.[1] In this section, we briefly set out the theories and previous empirical works. We also discuss the predictions that can be tested on our sample of local newspapers. This discussion motivates specifications in the empirical part, where we regress the changes in subscription prices, advertising prices and margins in 1990-1992 and 1994-1996 on a number of independent variables.

2.1 Switching Costs and Liquidity Constraints

In the models of Chevalier and Scharfstein (1996), Gottfries (1991), and Klemperer (1995), consumers incur a switching cost when changing supplier. In the short run, this will reduce the price sensitivity of the firm's customers. Effectively, this allows a firm to exploit captured customers by setting a high price to raise short-run profits. In the long run, a high price will induce consumers to search for other alternatives, and customers once lost are costly to win back. Clearly, when firms have access to capital markets, i.e. are able to borrow against future profits, they will not sacrifice long-run

[1] Most of the early empirical studies, e.g. Domowitz et al (1988), employed aggregate inter-industry data and found preciously little evidence of counter-cyclical tendencies. Rotemberg and Woodford

profits for short-run gains. However, liquidity constraints, for instance in recessions, could leave firms with little choice but to raise prices to meet debt repayments.

In their empirical application, Chevalier and Scharfstein (1996) used samples of supermarkets in the U.S. to test their hypothesis that liquidity-constrained firms tend to increase their prices, see also Chevalier (1995). They measured the importance of liquidity constraints by i) the market share of local chains compared to national chains in areas hit by a sharp drop in oil prices, ii) the market share of chains that had recently undertaken leveraged buy-outs[2] and iii) whether firms had undertaken leveraged buy-out. The evidence suggests that firms more likely to be liquidity constrained raised their prices in economic downturns. Compared to other firms, however, the difference in pricing is only in the range of a few percent. The results in Phillips (1995) also suggest that financial constraints often lead to price increases. However, Borenstein and Rose (1995) find that airlines, in or close to, bankruptcy lower their prices, but they find no effect on prices or sales of rival firms. Kennedy (2000) finds that there are adverse effects on the profits and sales of rival firms before a bankruptcy, but afterwards the rival firms' profits recover quickly.

Some features of the present data make a test of the predictions on liquidity constraints and customer markets attractive. First, as noted in the introduction, when Sweden fell into a recession in 1990, credits were tightened as a result of very large credit losses for all major banks. Using accounting information from newspaper firms, we can by their solvency, i.e. the ratio of own equity to total assets, broadly categorize them as being more or less liquidity constrained.[3] Firms with low solvency are likely to be most affected in the event of a tightening of credit markets. Solvency is measured one year prior to the period in question, which might be treated as exogenous, given that the recession was unexpected. Second, the revenue of a representative local newspaper roughly splits equal between the sales of subscriptions and advertising space. We argue that the average buyer of advertising space is much

(1999) give references to more recent work in this vein. Schmalensee (1989, p.987) takes as a stylized fact that "price-cost margins tend to be more strongly procyclical in more concentrated industries".

[2] There is an issue whether ii) is a valid test group. The reason is that leveraged buy-outs are not exogenous and a change in ownership may lead to substantial changes in firm strategy, not directly related to liquidity constraints.

[3] Studies of investment behavior have used profit margins, dividend payouts or differences in solvency, e.g. due to leveraged buy-outs, to measure the extent of liquidity constraints; for references see Hubbard (1998).

less attached to a particular newspaper, i.e. has lower switching costs, than the average subscriber. This is motivated by the ease with which advertisers can change source in response to a change in the cost of reaching consumers.[4] The reader, on the other hand, has grown accustomed to a newspaper's content and style, and is therefore unwilling to switch due to modest increases the subscription price. This conjecture is supported by studies of newspapers which usually find a highly inelastic relation between circulation and subscription price in the short-run; see references quoted in Lewis (1995).

Sales of subscriptions and advertising space are interrelated. As advertisers have a preference for newspapers with a large circulation, increasing the subscription price will involve a trade off between a higher subscription revenue and a loss of advertising revenue.[5] However, with circulation being insensitive to the subscription price in the short run, the newspaper has the option of increasing its total revenue by raising the subscription price. Given that this only has a minor effect on circulation, the advertising price may not need to be adjusted. This motivates our empirical specification, which treats the growth rate of the subscription price as independent of the growth rate of the advertising price.[6]

The prediction from the model of Chevalier and Scharfstein (1996) is that firms with the lowest solvency in 1989 should have the highest price increases for subscriptions in the 1990-1992 recession. It is not immediately clear that any prediction can be made for the 1994-1996 recovery. First, as the general economic conditions improved during this period, it gradually became easier to obtain new credits even for firms with low solvency. Second, it may be too costly for newspapers that have lost subscribers in a recession to win some back by lowering their prices. In

[4] Advertisers in national newspapers and on TV, usually producers of branded consumer products and large retail chains, are likely to have more long-term relationships that are costly to terminate, for instance due to quantity discounts. However, the advertising content of local newspapers is dominated by occasional promotions by local businesses, which suggests a reliance on short-term contracts.

[5] There is mixed evidence of the magnitude of the relation between subscription price and advertising sales. Thompson (1989) and Dertouzos and Trautman (1990), among others, have estimated systems of equations for cross section data. Using US data, Dertouzos and Trautman find some effects of subscription price on sales of advertising, with a price elasticity of -0.45. This is the long run effect however, since they use cross-section data for one year only. Using British and Irish data, Thompson finds no clear price-quantity relationship for either subscription or advertising price.

[6] In practice, of course, a newspaper will offer a large menu of advertising possibilities depending on the size and place of the ad, and often also sells subscriptions of different duration. The prices we use,

both periods, prices for advertising space should be essentially independent of the financial position.

2.2 Customer Flows

The demand functions might be more elastic during booms. If demand is more elastic in high demand states, prices, collusive as well as non-collusive, should be kept relatively low. One reason why demand is more elastic is that in high demand states, there is an inflow of new consumers, i.e. consumers who have not previously bought the product. If these consumers develop switching costs after their initial purchase, this induces firms to cut prices in high demand states in order to capture these new consumers for future exploitation; see Bils (1989) and Klemperer (1995). However, the argument rests on the assumption that the demand increase stems from an inflow of new customers, not higher demand from repeat purchasers, which would lead to higher prices.

Our data allows us to discriminate between high demand per consumer and a large inflow of new consumers. Although an upturn in the economy affects the demand per consumer in all local markets, certain areas will benefit relatively more. We measure this variation at the market level by the growth rate in the average per capita disposable income. To measure the importance of an inflow of consumers, we use statistics on migration patterns within the country. From this hypothesis, we expect to find lower subscription price increases in markets with a high inflow of consumers, measured as a percentage of the population. For the advertising market, on the other hand, large inflows would lead to higher demand for advertising from local business and would be more likely to raise than to lower advertising prices.

2.3 Implicit Collusion and Market Structure

It is often stated that the intensity of competition varies with demand conditions, and that firms have difficulties in sustaining implicitly collusive agreements when demand is unstable. Rotemberg and Saloner (1986) formalized the intuition in a model where demand fluctuates randomly between high and low states, but where firms can

annual full price subscription and price per millimeter for non-specified placement, are those on which people in the industry usually rely in their comparisons.

observe the realization of demand. The key intuition behind their result that price-cost margins may have a counter-cyclical tendency is that firms have a greater incentive to deviate when the value of future collusive profits is low, compared to the value of a current deviation. With no correlation in demand, it may be impossible to sustain fully collusive prices in periods with high demand. Empirically, this can be contrived as demand fluctuating randomly between high and low demand states; the issue of the temporal dependence of demand shocks has been addressed in several extensions of the model. Bagwell and Staiger (1997) consider the effects of randomness in the growth rates in demand.[7] Under the empirically plausible assumption of positive correlation of growth rates in demand, they show that price-cost margins tend to be pro-cyclical rather than counter-cyclical. Green and Porter (1984) model a situation where firms cannot observe the state of demand. This might lead to an equilibrium with temporal punishments following a low demand state, which could be observed as a pro-cyclical tendency of price-cost margins.

A number of studies have attempted to test predictions from the models of implicit collusion. One of the most comprehensive studies to date is Ellison's (1994) structural model of supply and demand in a railway shipping cartel in the 1880's, designed to test predictions from competing models of implicit collusion. He finds little of support for the mechanism in Rotemberg and Saloner (1986) but some evidence in favor of that in Green and Porter (1984). Borenstein and Shepard (1996) study the dynamics of prices and margins in regional retail gasoline markets in the U.S., using the fact that demand fluctuates over the year and that there is a delayed pass-through of some cost components. They find that predictable increases in future demand (cost) tend to increase (decrease) price-cost margins, which is consistent with models suggesting that implicitly collusive agreements are more difficult to sustain when the gains from deviation are large.

The bottom line is that the theories of implicit collusion are sensitive to the fine details of the game. While it is possible in some cases to argue that the assumptions of a specific model are satisfied, it is more common, as is done here, to treat them as unobservables. The most robust prediction of models based on implicit

[7] See also Kandori (1991) for the case with correlated demand shocks, and Haltiwanger and Harrington (1991) for the case with deterministic fluctuations in demand.

collusion is that the pattern of price adjustments should vary with the market structure, in the sense that monopoly firms do not need to take such considerations into account.[8] The data contains a number of newspapers that can be considered as pure monopolies, and there are also cases where a local newspaper has a very high market share and only faces weak competition from other local and/or national newspapers. Other markets are characterized by close competition between two newspapers. In the regressions, we use a dummy variable to capture the effects of market structure.

3 Data

The Swedish daily newspaper industry is well documented. Advertisers have an interest in knowing the circulation and geographical coverage of newspapers as well as prices for advertising. This information is provided by Tidningsstatistik AB. The Swedish association for newspaper publishers, Tidningsutgivarna, collects information on subscription rates. The government subsidizes some newspapers and a government agency, Presstödsnämnden, monitors the publishers' economic performance. In addition to these sources, we employ census data from Statistics Sweden (Statistiska Centralbyrån). The full data set contains information on all (133, as of 1992) newspapers in Sweden over the period 1975-1996.[9] Our analysis, covering the turbulent period 1990-1996, is restricted to local morning newspapers with three or more issues per week, which gives a sample of approximately 90 newspapers in each of the years.[10] Details of the variables are given in Appendix 1.

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[8] Prices and margins might only gradually adjust to new conditions and it is possible that the speed of the adjustment process depends on the market structure. Fisher and Konienczny (1995) find that monopoly newspapers adjust their prices more frequently, and by smaller nominal amounts than newspapers with one competitor. However, Thompson (1988) found no evidence of market structure or circulation being important determinants of the speed of adjustment of advertising prices. While it is conceivable that our market structure dummy also picks up differences in the speed of adjustment, we believe the two-year intervals we use to be sufficiently long for prices to be fully adjusted to new conditions.

[9] The balance sheet variables used to measure liquidity constraints are only recorded from 1989 to 1996, thus we are forced to restrict attention to the 1990-1996 period.

[10] We have excluded four evening newspapers, sold almost exclusively as single copies, since these only compete with local morning newspapers to a minor extent. The second group excluded is morning newspapers with coverage in a very large number of local markets. This group includes three large national morning newspapers, and two newspapers tied to religious and political organizations, which cover most of the country but have very low local market shares. Finally, we do not consider 37 newspapers with only one or two issues per week, read as a local or political additive to a regular

3.1 The Economy over the Sample Period

Most of the changes in cost and demand conditions, such as wages, interest rates, paper costs and general inflation affect newspapers irrespective of the local market in which they operate. We ignore changes in such general conditions as our analysis is aimed at explaining differences across newspapers. In the years prior to our sample period, GDP growth rates were high and unemployment low. The beginning of the 1990's saw a very severe recession, with a falling real GDP for three consecutive years, 1991-1993, and a dramatic increase in unemployment. Tightening of firm credits was a prominent feature of the recession. All major banks suffered considerable credit losses and had a large *ex ante* probability of bankruptcy, with the total sum of credit losses amounting to 5% of GDP in 1992. The government had to intervene to save a number of banks from bankruptcy. The total number of firm bankruptcies almost tripled from 1989 to 1992. This recession was followed by a recovery, although unemployment remained at a high level. The development of the economy is depicted in Figure 1. Naturally, there is some ambiguity as to when exactly the recession ended and the recovery began and we therefore split the sample into two sub-periods: 1990-1992 and 1994-1996. Prices are measured in January each year and for most newspapers, a price change takes place in January. The recession began in 1990 but was not anticipated in January that year.

[FIGURE 1 ABOUT HERE]

3.2 Market Level Data

Our definition of a market follows the standard Swedish municipal classification. We define a newspaper's home market as the municipality where it has its largest circulation.[11] The median firm has sixty-two percent of the total circulation in its home market. A newspaper's market level data is from its home market[12] and includes

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newspaper and viewed as distinctly different by government agencies and people involved in the industry. The exact numbers vary slightly over the years, the numbers above are for 1992.

[11] In one case only does the municipality where a firm has it its largest circulation change. Since this newspaper has roughly the same circulation in the two municipalities, we let its home market remain the same over the sample period.

[12] The economies in neighboring municipalities are likely to be closely connected. To simplify the empirical analysis, we assume that the economic conditions in all municipalities where a newspaper has circulation are perfectly correlated.

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demographic and income statistics. In the regressions we use the change in disposable income per capita over the periods expressed as a percentage,[13] denoted *INCGROWTH*. According to customer market theories, e.g. Bils (1989), the level of immigration affects the price level and thus, a change in the number of new consumers will cause a change in prices which we capture with the change in the ratio of immigration to total population, *IMMIGDIFF*.

3.3 Newspaper and Firm Level Data

To measure the development of nominal prices during the two periods, we use the annual subscription price and the price per millimeter for ads with a non-specified placement as of January 1.[14] The dependent variables are the growth rates of the two nominal prices *PSUBGROWTH* and *PADVGROWTH*, defined as 100*(P[t]-P[t-2])/P[t-2]. Note that it is irrelevant whether nominal prices or prices deflated by CPI are used, as general inflation is the same for all newspapers. The data also includes information from the income statement (e.g. total revenues and costs and goverment subsidies) and some balance sheet variables (e.g. own equity and total debt). Certain owners control several local newspapers[15] but for the vast majority, accounting data are broken down per newspaper, as most newspapers are individual firms. In the cases where no separate figures exist, we assume the accounting variables to be the same for all newspapers in the same firm. The income statement gives our next dependent variable, that is the change in price-cost margin, *MARGDIFF*, defined as 100*(PMC[t]-PMC[t-2]), where PMC is total revenues, including government subsidies, minus total costs, divided by total revenues. This measure is admittedly crude and is also the same for newspapers in the same firm, but it can give an

[13] Disposable income is aggregated over the year and is measured on December 31. It is an issue whether *INCGROWTH* should be measured as 100*(X[t]-X[t-2])/X[t-2] or 100*(X[t-1]-X[t-3])/X[t-3] for t=1992, 1996. In the regressions, we use the former definition, but this is not material to the results reported below. Similarly, we measure *IMMIGDIFF* by 100*(X[t]-X[t-2]).

[14] Subscription rates often remain unchanged for periods up to a year, potentially leading to a measurement problem in the dependent variable. However, most newspapers adjust prices in January such that this problem is likely to be of minor importance. Advertising rates are adjusted more frequently.

[15] As of 1994, some local newspaper had the same owner. One owner owned nine local newspapers, one owned seven, one owned five, two owned four and three owned three. In some cases, all newspapers are within the same firm, in other cases there are several separate firms. Matching the accounting data with circulations and subscription prices suggests that the major part of the firms' revenue stems from the industry, i.e. they seem to have very limited exposure to activities in other industries.

indication of whether a newspaper's short-run performance improved or deteriorated during a certain period.

We are interested in whether a newspaper's behavior depends on its market position. As noted above, some municipalities are the home market for only one local newspaper, while others have two local newspapers. There are no municipalities with more than two local newspapers with three or more issues per week. If there is no competing local[16] newspaper, the newspaper is referred to as a monopoly and the dummy variable *MONOPOLY* takes on the value one. Monopoly markets tend to be small in terms of population and, consequently, monopoly newspapers have a low total circulation. The competitive situation for newspapers with one local competitor may differ depending on their relative market shares, but no attempt is made here to make a finer distinction between newspapers in this category[17].

As the theory predicts that solvency should only play a role for newspapers facing a bankruptcy risk we use a dummy variable, *LOWSOLVENCY*, as a measure[18] of liquidity constraints.[19] The cutoff level is a solvency of 15 percent, but we have also experimented with values between 10 and 20.[20] For newspapers with one local competitor, there might be interaction in the pricing decisions. For instance, if a liquidity constrained newspaper raises its subscription price, it may also open the possibility for its rival to raise prices. This is captured by the variable

[16] In most of the monopoly markets, there is one or more non-local newspapers with a household coverage of more than three percent. As the home market accounts for 62.5 (62.3) percent of the total circulation for the average (median) newspaper, we expect there to be little strategic interaction between a local and a non-local newspaper.

[17] We have, however, tried to assign a dummy variable to each newspaper according to its position in the home market. There is considerable heterogeneity among newspapers in the duopoly markets. Some duopoly markets have two newspapers with roughly the same coverage. Furthermore, there are asymmetric duopoly markets where one of the newspapers has significantly larger coverage. Preliminary regressions showed no evidence of differences in the pricing pattern between different types of duopoly newspapers.

[18] As mentioned above, several newspapers may have the same owner. In one case, the owner, Nya Wermlandstidningen, owns a subsidiary whose solvency was low in 1989, and eight subsidiaries with high solvency in that year. We treat all newspapers with this owner as having high solvency. This is motivated by our intention to let low solvency be a measure of bankruptcy risk. If the owner has a strong financial position it could, if needed, make a capital infusion to save a subsidiary from bankruptcy.

[19] The reason for not using a continuous variable is that for a firm with high solvency, the bankruptcy risk is negligible, and it is irrelevant if solvency is e.g. 50 or 80 percent.

[20] If the cutoff level is set at 10 percent, 14 firms are in the *LOWSOLVENCY* category, at 15 percent there are 19 firms and at 20 percent, there are 27 firms.

RIVAL_LOWSOLVENCY, which takes the value one if the newspaper competes in the same market as a newspaper with solvency below 15 percent

The dummy variable *BANKRUPT* denotes four newspapers, all belonging to the same owner and having low solvency in 1989, in the sample that have actually become bankrupt in 1992. They are still present in the market and quote prices for this year. However, we do not have any financial data, such as margins, for these firms for 1992; thus they are missing in the *MARGDIFF* regression. Bankrupt newspapers may face different demand conditions due to consumer uncertainty regarding future publication and quality of the newspaper. These newspapers were reconstructed in 1993, but had a significantly lower circulation in the 1994-1996 period. The strategic interaction in the local market may also be affected by the bankruptcies. To control for the possibility of such effects, we introduce the dummy *RIVAL_BANKRUPT.*

[TABLE 1 ABOUT HERE]

As seen in Table 1, the sample contains 19 newspapers whose solvency was below 15 percent in 1989, four of which were bankrupt in 1992. The low solvency newspapers are over represented in the duopoly markets, 12/34 compared to 7/53. There is, however, no difference in the average size, measured as total circulation, between those with high and low solvency. Note that monopoly newspapers are, on average, smaller than duopoly newspapers. Market structure is largely exogenous, a smaller city cannot support more than one newspaper and larger cities tend to be duopolies.

[TABLE 2 ABOUT HERE]

Table 2 reveals that the average increases in the nominal price were higher in the first than in the second period, both in the subscription and the advertising market. Margins improved by, on average, 2.5 percent over the recession and declined by, on average, 5.5 percent in the recovery.[21] In the first period, it is obvious that firms that

[21] The differences in averages across periods can partly be attributed to growth in CPI and input costs. Growth in CPI, measured as of January, was 16 percent between 1990 and 1992, and 4 percent between 1994 and 1996. Hence, real prices increased by roughly 6 percentage points in each of the

eventually go bankrupt have much smaller price increases, 14 percent, than the average 22 percent. Other firms with low solvency exhibit subscription price increases that are three percent higher than the average, but advertising price increases that are close to average. A mirror image to that the group of firms with low solvency managed to increase their margins by 4.8 percentage points, or 2.2 percentage points more than the sample average. Interestingly, the growth rates of advertising prices show no important differences across newspaper types in the recession. For the second period, price changes are strikingly similar across groups of newspapers of different characteristics as well as for subscriptions and advertising markets. The only striking feature is a sharper fall in the margins in the monopoly group.

An important assumption in the analysis is that readers have switching costs and that circulation is relatively insensitive to price in the short run, such that it is possible to improve, temporarily, margins. We perform some simple checks to see whether this maintained assumption is satisfied for our sample. Although our data is insufficient to estimate demand functions for the newspapers, e.g. we lack measures of their quality, it is still possible to use information about circulation to provide some evidence. During the recession, average circulation fell by 1.5 percent only while nominal subscription prices rose by 22 percent, while CPI increased by 16 percent only. The correlation coefficient for growth in the subscription price and the change in circulation was -0.28. In the same period, the correlation between the change in margins and the growth in subscription price was 0.34. Altogether, this suggests that price sensitivity was indeed low, and that the newspapers increasing their subscription prices the most, were able to improve their margins substantially.

[TABLE 3 ABOUT HERE]

Table 3 gives the means and standard deviations of the independent variables. Most importantly, we see that there is very little variation in *INCGROWTH* and *IMMIGDIFF*. The fraction of firms categorized as *MONOPOLY* has risen from 60.3

periods. In the two periods, journalist wages increased by 7 and 3 percent, and the newsprint price declined by 5 percent in the recession and increased by 26 percent in the recovery. The improvements in margins in the recession and the decline in the recovery thus, to some extent seem to be caused by the movements in newsprint price.

percent in 1992 to 64.4 percent in 1994, due to the bankruptcies of three duopoly newspapers.

4 Results

Following the discussion above, our basic specification for the estimation of changes in prices is

$$\begin{aligned} Y_t &= \beta_0 + \beta_1 LOWSOLVENCY + \beta_2 RIVAL\_LOWSOLVENCY + \\ &+ \beta_3 BANKRUPT + \beta_4 RIVAL\_BANKRUPT + \\ &+ \beta_5 IMMIGDIFF + \beta_6 INCGROWTH + \\ &+ \beta_7 MONOPOLY + \varepsilon \end{aligned}$$

where *Y=PSUBGROWTH, PADVGROWTH* and *MARGDIFF*, *t*=1992,1996, is the end of the period for which the price changes are estimated and ε is an error term.[22]

The results for the two periods are shown in Tables 4a and 4b.[23] Newspapers in bankruptcy do not report any financial variables for 1992. Thus, we have no observations for changes in margins for these firms and *BANKRUPT* is excluded as an explanatory variable. No firms went bankrupt in the second period, hence the variables *BANKRUPT* and *RIVAL_BANKRUPT* are excluded in the second set of regressions. The variables included can help explain the development of prices and margins for the 1990-1992 period.[24] However, the adjusted R-square for the 1994-1996 period is practically zero. Hence, only in the sharp downturn are there any differences across newspapers in their pricing behavior and the development of margins. Stated differently, in the recovery period, price changes follow from idiosyncratic changes in cost or demand conditions. In the following, we therefore focus our discussion to the former period.

[22] The raw correlation between the rates of change of the subscription and advertising price is 0.15 and 0.01 for the 1990-1992 and 1994-1996 period, respectively. The correlation between the error terms in the subscription price and the advertising price regressions are 0.02 and 0.05. This supports our specification that treats the rates of price changes in the two markets as nonrelated.

[23] The difference in the number of observations in Table 4a is due to missing accounting data for the firms for 1992; hence it is not possible to compute the margin. In four cases, the lack of accounting data is due to bankruptcy. In Table 4b, the difference in number of observations is due to missing accounting information for one newspaper.

[TABLES 4a AND 4b ABOUT HERE]

4.1 Switching Costs and Liquidity Constraint

Newspapers with low solvency raised their subscription prices significantly more than did others. In nominal terms, the difference is 3.6 percentage points, which is 17 percent more than at the means (21.7 percentage points). In contrast, we found no difference in pricing between the financially constrained firms and those with good solvency in the advertising market.[25] The margins of the financially constrained firms rose by 2.5 percentage points more than those of the others. This is a sharp improvement, which should be attributed to the higher subscription prices, as there was no significant difference in the development of the advertising price. Hence, we find evidence that firms with a high default risk can, and do, exploit their customer base in an attempt to raise short-run profits. Firms competing with liquidity constrained firms also exhibit significantly larger increases in prices. The main results from the three regressions on the 1990-1992 data conform to those reported in Chevalier and Scharfstein (1996).

The newspapers that went bankrupt in 1992 had much smaller price increases than the average.[26] There may be two explanations to this finding. First, if it becomes clear to readers that a newspaper will inevitably become bankrupt in the near future, then the demand for subscriptions will drop. In order to upkeep its circulation, the newspaper may be forced to lower its subscription prices. Second, these firms may have gone bankrupt just because they did not raise prices sufficiently. A closer examination of price changes for the periods 1990-1991 and 1991-1992 suggest that it is in the latter year that the bankrupt newspapers behave differently. Between January

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[24] Using 1991-1993 to capture the downturn does not change the results qualitatively, using the period 1989-1991 is not possible due to the lack of data on solvency for 1988.

[25] Using 10 percent as the cutoff level for *LOWSOLVENCY* changes the estimates for *LOWSOLVENCY* to 2.48**(1.26) and 2.20(1.68) and for *RIVAL_LOWSOLVENCY* to 1.47(0.0239) and 2.96*(1.80), in Table 4a, columns (1) and (3), respectively.
Using 20 percent as the cutoff level changes the estimates for *LOWSOLVENCY* to 3.00** (1.20) and 2.05 (1.37) and for *RIVAL_LOWSOLVENCY* to 2.53 (1.93) and 3.50*(1.88), in Table 4a, column (1) and (3) respectively. Estimates in Table 4b are essentially unchanged.

[26] The estimates in the *PSUBGROWTH* regression for 1990-1992 do not change a great deal if *BANKRUPT* is excluded; the point estimate for *LOWSOLVENCY* changes to 3.63***(1.29). For the *PADVGROWTH* regression for 1990-1992, adjusted R-square almost falls to zero, which is not surprising since *BANKRUPT* is the only significant variable in that regression.

1 1991 and January 1 1992, the bankrupt newspapers increased their subscription prices by only 3.6 percent compared to 10.7 for the sample average, and 12.6 for those with low solvency that avoided bankruptcy. In the prior year, there were no pronounced differences between the pricing behavior of different types of newspapers, although the ones with low solvency raised their prices by 2 percent more than the sample average. For the full sample, circulation fell by 1 percent, on average, in each of the years 1990-1991 and 1991-1992. The circulation for the newspapers that went bankrupt in 1992 fell by 14 percent and 15 percent in each of the two years, respectively. Altogether, this indicates that readers may have realized already in 1990-1991 that some newspapers would not survive, and that these were therefore forced to limit price increases in 1991-1992.

4.2 Customers Flows

Markets with large inflows of new consumers should exhibit lower prices according to customer market theories. A change in consumer inflows will then cause a change in the price. However, we find no evidence of such an effect.[27] One reason may be the newspapers' ability to directly target new subscribers with discounted introductory offers, without having to cut their regular subscription prices. Another explanation is that the number of new consumers, compared to the total, is too small to make it profitable to cut subscription prices in order to capture them.[28] Contrary to our expectations, the estimated coefficient on *INCGROWTH* is negative in both the *PSUBGROWTH* and *MARGDIFF* regression. The likely reason for this is that small local differences in the economic conditions make it difficult to trace any effect.

4.3 Implicit Collusion and Market Structure

We find no significant differences in either subscription price or advertising price, or in price cost margins, between firms facing competition and monopolies in any of our regressions. For subscription prices these findings are not surprising, as any newspaper can rapidly detect the moves of its rivals, which is contrary to a basic assumption in the model. Whether the same is true for advertising prices is unclear.

[27] Further testing of the prediction on other subsamples from the full data set from 1975 to 1996, has not provided any more substantive support.

The prices in our data are list prices per millimeter of non-specified placement, but the possibility remains that some large advertisers get secret discounts. According to people in the industry, this is quite common for the period studied.[29]

5 Conclusions

In this paper, we have examined the behavior of prices and margins in the Swedish newspaper industry during a sharp economic downturn and a following recovery. The purpose was to evaluate some mechanisms that could lead to a countercyclical tendency of prices and margins, such that prices are relatively high in periods of low demand. In particular, we focused on the effects of liquidity constraints in a recession that more than tripled the number of bankruptcies in the economy.

We find that newspapers with low solvency at the beginning of the recession increased their subscription prices and that their margins improved, relative to other newspapers. Our interpretation is that their low solvency implied a bankruptcy risk, which could only be avoided by exploiting readers whose switching costs were high. In the same period, increases in advertising prices did not depend on financial strength, consistent with the notion that buyers of advertising space have great freedom to switch supplier in response to a price increase. In the recovery, when credit conditions gradually eased, measures of solvency could not explain any of the variation in pricing behavior. These results are consistent with the predictions of Chevalier and Scharfstein (1996). We also examined whether pricing patterns differed between monopoly newspapers and duopoly newspapers, as might be expected if the possibility to sustain implicitly collusive agreements depends on the state of demand, see e.g. Rotemberg and Saloner, 1986, and Green and Porter, 1984. However, we find no evidence for this hypothesis. Neither were the pricing pattern dependent on regional differences in migratory patterns and income growth, as suggested by models along the lines of Bils (1989).

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[28] In our companion paper, Asplund, Eriksson and Strand (2001), we show that newspapers in areas with a high proportion of new consumers have a higher proportion of subscriptions sold at a rebate.

[29] Interview with Madelaine Skedung, Dagens Nyheter AB.

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Table 1. Descriptive statistics by firm type as of 1992, regression sample.

	# (1992)	Mean Circulation	#LOW-SOLVENCY	#RIVAL_LOW-SOLVENCY	#BANK-RUPT	#RIVAL_-BANK-RUPT
All newspapers	87	27600 (33200)	19	11	4	5
LOWSOLVENCY=1 (Solvency<15%)	19	27500 (20300)	19	2	0	0
BANKRUPT=1	4	13000 (4400)	4	0	4	0
BANKRUPT=0	15	33500 (67700)	15	2	0	0
LOWSOLVENCY=0 (Solvency>15%)	68	26700 (20500)	0	9	0	5
MONOPOLY=1	53	20400 (15200)	7	0	0	0
MONOPOLY=0	34	39100 (48200)	12	11	4	5

Standard errors are in parenthesis.

Table 2. Descriptive statistics by firm type as of 1992, regression samples

	PSUBGROWTH		*PADVGROWTH*		*MARGDIFF*	
	1990-92	1994-96	1990-92	1994-96	1990-92	1994-96
All newspapers	22.5 (5.59)	9.94 (6.55)	16.5 (4.07)	9.77 (4.30)	2.58 (5.45)	-5.57 (6.55)
LOWSOLVENCY=1 (Solvency<15%)	23.4 (6.25)	10.2 (7.67)	14.6 (4.92)	8.78 (5.05)	4.79 (5.42)	-5.26 (8.43)
BANKRUPT=1	14.3 (3.28)		8.87 (6.96)			
BANKRUPT=0	25.8 (4.27)		16.1 (2.97)		4.79 (5.42)	
LOWSOLVENCY=0 (Solvency>15%)	21.8 (5.61)	9.88 (4.17)	16.6 (4.28)	10.0 (4.13)	2.07 (5.37)	-5.64 (6.11)
MONOPOLY=1	22.0 (5.26)	9.53 (4.23)	17.1 (3.10)	9.89 (3.37)	2.01 (5.56)	-6.38 (5.78)
MONOPOLY=0	23.2 (5.99)	10.6 (6.02)	15.5 (5.15)	9.56 (5.69)	3.53 (5.22)	-4.08 (7.67)

Standard errors are in parenthesis.

Table 3. Descriptive statistics, independent variables 1990-1992, 1994-1996.

	1990-1992	1994-1996
LOWSOLVENCY	0. 218	0.184
	(0. 415)	(0.389)
RIVAL_LOWSOLVENCY	0.126	0.0805
	(0.334)	(0.273)
BANKRUPT	0.0460	
	(0.211)	
RIVAL_BANKRUPT	0.0575	
	(0.234)	
IMMIGDIFF	-0.489	-0.230
	(0.416)	(0.405)
INCGROWTH	18.1	2.19
	(1.38)	(1.02)
MONOPOLY	0.602	0.644
	(0.491)	(0.482)

Standard errors are in parenthesis.

Table 4a. Regression Results, Prices and Margins 1990-92

Variable	*PSUBGROWTH*	*PADVGROWTH*	*MARGDIFF*
CONSTANT	27.2***	15.9***	22.6**
	(7.68)	(4.72)	(9.13)
LOWSOLVENCY	3.61***	-0.725	3.47**
	(1.30)	(1.09)	(1.57)
RIVAL_LOWSOLVENCY	4.13**	0.209	3.18
	(2.03)	(1.87)	(1.97)
BANKRUPT	-11.5***	-7.03**	
	(2.00)	(3.35)	
RIVAL_BANKRUPT	-0.215	-0.00171	1.94
	(2.11)	(2.20)	(2.6)
IMMIGDIFF	0.607	0.781	0.947
	(1.46)	(0.960)	(2.09)
INCGROWTH	-0.298	0.103	-1.15**
	(0.421)	(0.265)	(0.481)
MONOPOLY	-0.550	1.01	-0.404
	(0.67)	(1.51)	(1.89)
Adj R^2	0.168	0.117	0.069
Number of obs.	87	87	81

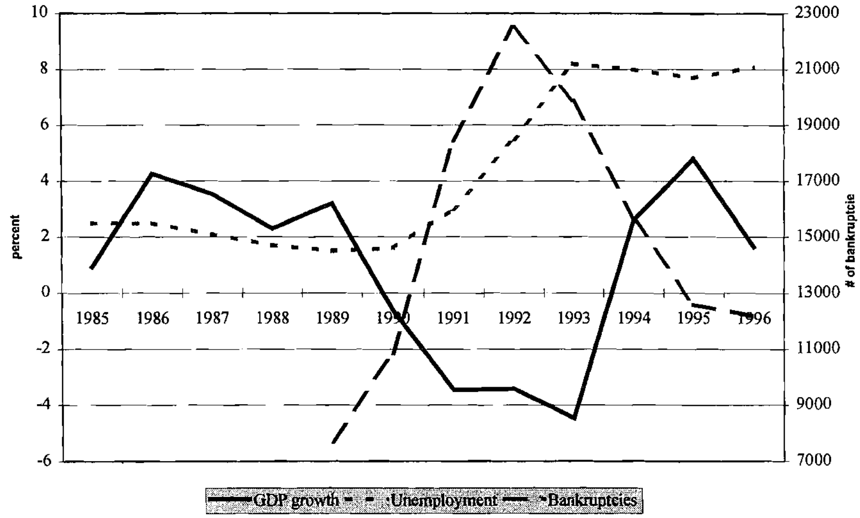
Standard errors, using White's robust covariance matrix, are in parenthesis.
Variables starred *** are significant at the 1% level, with ** at the 5% level and with * at the 10% level.

Table 4b. Regression Results, Prices and Margins 1994-96

Variable	*PSUBGROWTH*	*PADVGROWTH*	*MARGDIFF*
CONSTANT	3.31**	9.61***	-5.30**
	(1.49)	(1.65)	(2.26)
LOWSOLVENCY	-0.170	1.32	-0.249
	(1.61)	(1.32)	(2.12)
RIVAL_LOWSOLVENCY	-0.664	-2.85	-4.06
	(2.19)	(2.07)	(2.79)
IMMIGDIFF	-2.06	1.07	-1.31
	(1.35)	(0.889)	(1.35)
INCGROWTH	-0.139	-0.236	-0.490
	(0.390)	(0.385)	(0.712)
MONOPOLY	-1.21	-0.263	-2.95
	(1.21)	(1.32)	(1.81)
Adj R^2	-0.011	-0.007	0.006
Number of obs.	87	87	86

Standard errors, using White's robust covariance matrix, are in parenthesis.
Variables starred *** are significant at the 1% level, with ** at the 5% level and with * at the 10% level.

Figure 1. The Swedish economy 1985-1996.



Chapter 5

Essay IV: Statistical Discrimination and Sex Stereotypes in the Labor Market

Statistical Discrimination and Sex Stereotypes in the Labor Market*

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Abstract

Time out for child care is unevenly distributed between the sexes. This paper investigates if this can be explained by self-fulfilling sex stereotypes in the labor market. The paper provides a model of distribution of time out for child care based on statistical discrimination and human capital investments. The model has three equilibria. In one equilibrium, time out for child care evenly distributed between the sexes. In the second equilibrium, there is full specialization. The third equilibrium is an intermediate case, where time out for child care is unevenly distributed without full specialization. There are no differences in ability or variance of ability between the sexes, the only differences between the equilibria are the self-fulfilling expectations of firms and workers. (*JEL* D63, J71)

Keywords: Labor market discrimination, statistical discrimination.

1. Introduction

The effects of parental leave policies have been the subject of many studies. The regulations of parental leave benefits vary considerably between countries and

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over time. A large number of papers have utilized this variation for empirical examinations of the effects of different reimbursement rates, the time and the qualification requirements to be eligible for benefits, e.g. minimum period of work before qualifying for the benefit. One of the issues that has been addressed is the effects on equality between the sexes. Many empirical studies have found that generous parental leave benefits increase the female labor participation rate. On the other hand, there are some concerns that generous parental leave benefits may have a negative effect on wages and career opportunities for women. Almost all of the parental leave benefits are used by women and time out for child care is likely to impose some costs on employers. Although the employer does not usually the parental leave benefits, she may incur costs for recruiting and training a replacement. These costs may differ between assignments. When the costs are substantial wages for women should be lower, or the employer should refuse to hire women, given that women are more likely to take time out for child care.

Parental leave benefits are usually exclusively given to the mother or distributed to both parents according to their choice. In Sweden one month is reserved for each parent since 1994, however. One reason for this is to give the child better contact with both parents, but increased equality between the sexes in the labor market has also been put forward as an argument. Two effects are often mentioned in this context. The unequal distribution of parental leave may create discrimination against women, and some men are subject to pressure from employers and co-workers not to use their parental leave. The latter may be explained by male parental leave being unexpected and signaling a lower commitment from the worker.

There has been an intensive public debate in Sweden on the pros and cons of restricting the free distribution of the parental leave between the parents. The obvious argument against the restriction is that it gives the parents a smaller choice set. The counter argument is that, these individually rational choices do not take into account that they may lead to expectations of women being mainly responsible for child care. This paper will show that such expectations might be self-fulfilling, so that women are expected to stay home for a long period of time with their children and employers therefore assign them less qualified and/or less well paid assignments. Hence, when parents decide on the distribution of parental leave, men are likely to have jobs with higher wages, which makes absences more costly.

There is a literature on self-fulfilling racial stereotypes, where the main idea is that there may be multiple equilibria of employers' expectations and workers' human capital investments. If employers' expectations are low, a discriminated group may invest less, since they are less likely to be assigned a job where their human capital investment is rewarded. Employers then find that the discriminated group has actually invested less in human capital. All employers and all members of the discriminated group act rationally, but there may be pareto superior equilibria, where employers have high expectations and workers in the discriminated group have high human capital investments.

One difference between racial discrimination and sex discrimination due to differences in time out for child care, is that higher expectations on one sex lead to lower expectations on the other, given that the total amount of parental leave does not change. One similarity between this model and some models of racial discrimination is that returns to human capital investments is different for different groups.

In this model, the choice of the distribution of parental leave benefits between the parents is modelled as a function of preferences for parental leave and the parents' wages. If men and women have different wages, they balance the gains from letting the high wage earner specialize in labor market work against the disutility of an unequal distribution of home and labor market work. Productivity varies between individuals in the model, and is not perfectly observable. Employers, who assign individuals to different jobs, observe workers' productivity and sex, preferring workers with high productivity and low absence.

Certain aspects of parental leave policies, which are certainly important for a normative analysis, are not modelled in this paper; for example, as mentioned above a more even distribution of the parental leave may affect the children. In this analysis, the costs for sex stereotypes are limited to the disutility of an unequal distribution of time out for child care between parents. However, it is possible that a change in sex stereotypes in the labor market will also affect other aspects of life in important ways. In this paper, preferences for parental leave are exogenous, but this may not be the case in reality. Changes in gender roles may also affect preferences for parental leave.

The contribution of this paper is that it shows that there might exist multiple equilibria, where time out for child care and labor market participation are either equally or unequally distributed between the sexes, and that this multiplicity of equilibria can be derived without any difference in average ability, distribution of ability or observability of ability between the sexes. The outline of the paper is as follows. In section 2, we discuss the literatures on parental leave policies and discrimination. In section 3, the model is derived and section 4 concludes.

2. Literature Review

The theoretical literature on racial discrimination is, to a large extent relevant, for sex discrimination. Sex discrimination and racial discrimination are similar in that sex and race are easily observed by employers. The most important difference is that racial stereotypes are noncorrelated for different races, higher expectations on one race do not affect expectations on others. This is not the case with sex stereotypes as the supply of wage labor and household work for men and women are interdependent decisions. If one spouse works more, the labor supply of the other will be affected; an effect that should be accounted for in a model of sex discrimination.

The two main strands of literature on discrimination are a theory of taste based discrimination, following Becker (1957) and a theory of statistical discrimination, following Phelps (1972) and Arrow (1973). The two theories are not necessarily conflicting. One argument often put forward against the taste based theory is that competition among employers will reduce discrimination if employers differ in taste. The statistical theory developed by Phelps relies on noisy observations of productivity. An employer will put more weight on the signal if it is less noisy. Phelps assumes the signal to be more noisy for minority groups. This assumption is also crucial in Lundberg and Startz (1983) who study the effects of affirmative action policies in a statistical discrimination model. Productivity is endogenously determined in their model, since human capital investments are affected by the wage, which is affected by expectations on productivity. Coate and Loury (1993) provide a model with self-fulfilling negative stereotypes without any group differences.

There is an empirical literature studying the effects of different parental leave policies. Rumh (1998) compares the effects of different parental leave policies in Europe. He finds that extended parental leave increases the labor participation rate for women, but reduces their relative wages, at least for long periods of parental leaves. The increase in labor participation is explained by the fact that women must work to be eligible for parental leave benefits. The reductional effect on wages may be explained by replacement and training costs during time out for child care. Stafford and Sundström (1996) study the effects of time out for child care for men and women and find that men lose three times as much as women for a year out. One natural interpretation is that this is due to signaling effects and that time out is unexpected for men.

3. The Model

In this model perfectly competitive firms set wages according to the expected productivity of workers. This productivity is not perfectly observed, so the firms set a wage scheme where wages depend on expected productivity given an imperfect signal of the productivity. Workers make two decisions; how much to invest in human capital and how to divide time out for child care. Human capital increases both productivity and the realization of the signal of productivity observed by the firms. It also raises the wage, since the wage depends on the signal.

All workers belong to a household consisting of a male and a female worker, all workers have one unit of time which can be used for labor or household work. All households must devote one unit of time to household work[1] and they decide how they want to divide household and labor market activities between the man and the woman. The decisions on human capital investments and division of work within the household are affected by the wage schemes set by the firms.

An equilibrium is characterized by that the firms' expectations on human capital investments and the division of work within the household, are fulfilled by the workers' behavior. It turns out that there might be both symmetric (where both sexes behave in the same way) and asymmetric (where one sex works more and invests more in human capital) equilibria.

In this section, we will first model the firms' wage schemes, the workers' human capital investments and the division of work within the household. Then, we will solve for the equilibria and examine welfare.

3.1. Wages, Human Capital Investments and Division of Labor

Wages, human capital investments and division of labor between the sexes are interdependent decisions. In the sections below, we will solve for each decision separately, taking the other decisions as given. The timing is assumed to be as follows. First, some expectations on labor division among the sexes are given. Given these expectations, employers offer wage schemes, which the workers take into account when investing in human capital. Finally, households decide on how to divide labor between men and women. In equilibrium, expectations on labor division are self-fulfilling.

[1]The assumption that households must devote half their time to household work simplifies some expressions. If the households have to devote some other fraction of time to household work, the results below will not change qualitatively.

3.1.1. The Wage Scheme

Each individual has one unit of time that can be used either for work or for child care. Perfectly competitive firms set wages equal to the expected productivity of the workers. The productivity is not perfectly observable to the firms; only a test score with a normally distributed error term is observed. As will be shown below, the expected productivity of an individual worker is a weighted average of the worker's test score and the average productivity of workers of his/her sex. The productivity MP is given by the following expression:

$$MP_i^s = a_i + bX_i^s(1 - d(1 - l_i^s)).$$

Subindex i denotes individual i and superindex s denotes sex, where an individual is either male (indexed by m below) or female (indexed by f below). The individual productivity parameter, a_i, is normally distributed, with the same distribution for both sexes. X_i^s denotes the human capital of individual i of sex s. In equilibrium all individuals of the same sex will invest the same amount of human capital, but investments may differ between sexes. There is a constant for the return to human capital investments, determined by the parameter b. The amount of work by one individual is l_i^s and since each individual is endowed with one unit of time out for child care is $(1 - l_i^s)$. The return to human capital is assumed to be negatively affected by absence from work, which might be due to costs for training replacements, depreciation of human capital during time out for child care or costs for interruption of production during absences from work. The term $(1 - d(1 - l_i^s))$ captures this effect. The costs for the employer is assumed to be proportional to the time out for child care, so that d is a constant.

Workers can invest in human capital at a cost C, given by

$$C = 0.5cX^2. \tag{3.1}$$

The marginal productivity is not perfectly observable. A test score T ,consisting of marginal productivity and a normally distributed error term, ε, is observed. The realization of ε is known by the worker, but not by the employer its distribution is the same for both sexes.

$$T_i^s = MP_i^s + \varepsilon_i^s \tag{3.2}$$

Workers are matched with each other according to their ability on the marriage market, so that both workers within a couple have the same a_i and ε_i.[2] The optimal wage contract with exogenous labor supply is (see Lundberg and Startz 1983)

$$w_i^s = \overline{MP}^s + B(T_i^s - \overline{T}^s)$$

where B is a constant between 0 and 1. The means of MP and T is denoted $\overline{MP}$ and $\overline{T}$, respectively. B is determined by the variance of ability, σ_a^2, and the variance of the error term, σ_ε^2. If the variance of the error term is high, the result of the test score will provide less information and less weight will be put on the test score, i.e. B will be low. See the Appendix for proof.

It is easy to see that human capital will be socially underprovided, since only a fraction B of the return of human capital investments will be visible in higher wages.

Inserting the expressions for $\overline{MP}$, T_i and $\overline{T}$ into the expression for the wage gives the wage for men

$$w_i^m = (1-B)(\overline{a} + b(1-d(1-\overline{l^m}))\overline{X^m}) + B(a_i + b(1-d(1-l_i^m))X_i^m) + \varepsilon_i)$$

and the wage for women

$$w_i^f = (1-B)(\overline{a} + b(1-d(1-\overline{l^f}))\overline{X^f}) + B(a_i + b(1-d(1-l_i^f))X_i^f) + \varepsilon_i)$$

These expressions can be simplified by utilizing that $X_i^m = \overline{X^m}$, $X_i^f = \overline{X^f}$, $l_i^m = \overline{l^m}$, $l_i^f = \overline{l^f}$ in equilibrium for all i. See the Appendix for proof.

$$w_i^m = (1-B)\overline{a} + Ba_i + b(1-d(1-l_i^m))X_i^m + B\varepsilon_i$$

$$w_i^f = (1-B)\overline{a} + Ba_i + b(1-d(1-l_i^f))X_i^f + B\varepsilon_i$$

Let us collect the individ specific terms and define

$$m_i = (1-B)\overline{a} + Ba_i + B\varepsilon_i.$$

We obtain

$$w_i^m = m_i + b(1-d(1-l_i^m))X_i^m \tag{3.3}$$

[2] Another way of specifying the model would be to randomly match workers on the marriage market. The motivation for the specification used here is that there is a positive relationship between the income of spouses.

$$w_i^f = m_i + b(1 - d(1 - l_i^f))X_i^f \tag{3.4}$$

as men and women have the same realization of a_i and ε_i within each couple.

[FIGURE 1 ABOUT HERE]

Figure 1 depicts the wages for men and women. Note that the difference in wages between the spouses is the same for all couples, as they are matched according to m_i.

3.1.2. Labor Division and Human Capital Investments

The household must make two decisions: First, how much each individual should invest in human capital and second, how to divide time out for child care. We derive expressions for these decisions in order to solve for the equilibria. The total amount of time available for work and child care is exogenous and set to unity for all individuals, so that the total amount of time the household is endowed with is two units one of which must be allocated to child care. The household's utility function is

$$U_i = w_i^m l_i^m + w_i^f l_i^f - 0.5c(X_i^m)^2 - 0.5c(X_i^f)^2 - k(l_i^m - l_i^f)^2,$$

where the last term is the cost for deviation from an equal division of parental leave. The quadratic functional form captures the idea of increasing marginal disutility for deviating from an equal division of household and wage labor and k is a constant. The quadratic form of the cost function for acquiring human capital reflects increasing marginal costs for human capital investments and c is a constant. For a given investment in human capital, we obtain the following first-order condition for labor supply (and, implicitly, the parental leave decision):

$$l_i^m = \frac{1}{2} + \frac{w_i^m - w_i^f}{8k} \tag{3.5}$$

$$l_i^f = \frac{1}{2} + \frac{w_i^f - w_i^m}{8k}. \tag{3.6}$$

Note that l_i^m and l_i^f are independent of $a_i, \overline{a}$ and ε_i, as the difference in wages is the same for all couples.

Noting that w_i^m and w_i^f depend on human capital investments according to (3.3) and (3.4), and taking the first order condition for X_i^m and X_i^f give

$$X_i^m = \frac{l_i^m Bb(1 - d(1 - l_i^m))}{c} \tag{3.7}$$

and

$$X_i^f = \frac{l_i^f Bb(1 - d(1 - l_i^f))}{c}. \tag{3.8}$$

3.2. Derivation of Equilibria and Comparative Statics

It is possible to obtain analytical solutions for the system of equations (3.3) to (3.8), in terms of the parameters in the model. Solving for l_i^m, we get three solutions (assuming that the restriction $0 \leq l_i^m \leq 1$ is satisfied). One solution, the symmetric equilibrium, is that labor is equally divided between the spouses. The second solution, the full specialization equilibrium, is that one sex specializes in household work and the other in labor market work. In the third solution, the partial specialization equilibrium, one sex works more in the labor market and the other does more household work, but both sexes do some labor market and some household work.[3] The solutions turn out to be independent of the individ-specific parameters a_i and ε_i, so that l_i^m, l_i^f, X_i^m and X_i^f are equal for all i. From now on the sub-indexes i for individuals are therefor skipped for these variables; see the Appendix for proof.

3.2.1. The Symmetric Equilibrium

The symmetric equilibrium, where $l_m = l_f = \frac{1}{2}$, exists for all parameter configurations. Solving the equation system (3.3) to (3.8) yields

$$X^m = X^f = \frac{Bb(1 - \frac{d}{2})}{2c},$$

$$w_i^m = w_i^f = m_i + b(1 - \frac{d}{2})(\frac{Bb(1 - \frac{d}{2})}{2c})^2.$$

[3] Actually, there are two solutions of the full specialization type and two of the partial specialization type. Men do more labor market work in one of these and women in the other. We can concentrate the analysis on the case of $l^m \geq \frac{1}{2}$, without loss of generality. The analysis is fully analogous in both cases, as there are no differences between sexes in the assumptions of the model .

Evaluating the utility function yields

$$\begin{aligned} U_i^S &= w_i^m l^m + w_i^f l^f - 0.5c(X^m)^2 - 0.5c(X^f)^2 - k(l^m - l^f)^2 = \\ & m_i + b(1-\frac{d}{2})(\frac{Bb(1-\frac{d}{2})}{2c}) - c(\frac{Bb(1-\frac{d}{2})}{2c})^2, \end{aligned}$$

where the superscript S denotes the full specialization equilibrium.

3.2.2. The Full Specialization Equilibrium

In the full specialization equilibrium wages are higher for individuals of the sex that is expected to work on the labor market. If the valuation of an equal division of household and labor market work is high enough, i.e. k is high enough, households will not fully specialize. The derivation of the conditions under which the full specialization equilibrium exist follow below. Solving for human capital investments, wages and utility, we get:

$$\begin{aligned} X^m &= \frac{bB}{c} \\ X^f &= 0 \\ w_i^m &= m_i + b\frac{bB}{c} \\ w_i^f &= m_i \\ U_i^F &= m_i + b\frac{bB}{c} - \frac{c}{2}(\frac{bB}{c})^2 - k, \end{aligned}$$

where superscript F denotes the full specialization equilibrium.

We now turn to the conditions for existence of the full specialization equilibrium. If a household changes the division of labor, it will lose the difference in wage between the sexes, $b\frac{bB}{c}$. The increase in utility from a more equal division of labor on the margin is the derivative of disutility of an unequal division of labor $\frac{dk(l^m-l^f)^2}{dl^m} = \frac{dk(2l^m-1)^2}{dl^m} = 4k$, evaluated at $l^m = 1$. Hence, the parameters of the model will have to fulfill the following expression for the fully specialized equilibrium to exist:

$$k \leq \frac{Bb^2}{4c}$$

3.2.3. The Partial Specialization Equilibrium

In the partial specialization equilibrium one sex specializes in the labor market work and one in household work, but the specialization is not complete. We obtain more complicated expressions than for the other equilibria,

$$l^m = \frac{1}{2} + \frac{\sqrt{(b^2Bd^2 + 16kc - 4Bb^2)}}{2b\sqrt{B}d} \tag{3.9}$$

$$l^f = \frac{1}{2} - \frac{\sqrt{(b^2Bd^2 + 16kc - 4Bb^2)}}{2b\sqrt{B}d} \tag{3.10}$$

$$X^m = \frac{l^m Bb(1 - d(1 - l^m))}{c}$$

$$X^f = \frac{l^f Bb(1 - d(1 - l^f))}{c}$$

$$w_i^m = m_i + b(1 - d(1 - l_i^m))X_i^m$$

$$w_i^f = m_i + b(1 - d(1 - l_i^f))X_i^f$$

$$U_i^P = m_i + \frac{(2 - B)(d - 2)dB^2b^4 + \left(4d - \frac{1}{2}d^2 - 2(1 + Bd - B)\right) 4kcBb^2 + 2(1 - B)16k^2c^2}{2b^2Bd^2c},$$

where superscript P in the last equation denotes the partial specialization equilibrium. This equilibrium does not exist for all parameter configurations. For some configurations, l^m is above one, which is inconsistent with the model, and for some configurations the expression under the root is negative. The following parameter configurations satisfy both restrictions:

$$\frac{Bb^2}{4c} - \frac{b^2Bd^2}{16c} \leq k \leq \frac{Bb^2}{4c}.$$

[FIGURE 2 ABOUT HERE]

Figure 2 depicts the existence of equilibria. In the analysis above, we have concentrated on the cases where $l^m \geq \frac{1}{2}$. Note that l^m increases in k for the partial specialization equilibrium, that is, the degree of specialization increases in the valuation of an equal division of household and labor market work. This may seem unintuitive. In equilibrium, the marginal disutility of increases in inequality must equal the marginal gains from higher income for the household, i.e. $w^m - w^f$. Wages are determined by expected productivity, and the difference in productivity increases with the difference between l^m and l^f. Hence, an increase in k must be accompanied by an increase in l^m in the partial specialization equilibrium. However, for large k, only the symmetric equilibrium exists.

3.2.4. Welfare Analysis

Given parameter values for which more than one equilibrium exist and assuming that the return to human capital investments is positive, it can be shown that utility is highest in the full specialization equilibrium, second highest in the partial specialization equilibrium, and lowest in the symmetric equilibrium. However, as noted in the introduction, it is reasonable to assume that there are factors outside the model that should be considered before making any policy recommendations. The restrictions on the parameter d in the propositions below, is motivated by that the return to human capital investments becomes negative for very high d.

Proposition 3.1. *For parameter values where the partial specialization equilibrium exists, $\frac{Bb^2}{4c} - \frac{b^2Bd^2}{16c} \leq k \leq \frac{Bb^2}{4c}$ and where the return to human capital investments is positive for all l^m, $d \leq 2$, the utility in the different equilibria is ordered as follows: $U_i^S \leq U_i^P \leq U_i^F$*

Proof

First we prove that $U_i^F \geq U_i^P$

$$U_i^F - U_i^P = \left(Bb^2 - 4kc\right) \frac{(2-B)b^2Bd + 4(1-B)kc}{b^2Bd^2c} \tag{3.11}$$

The second part of (3.11) is positive, since $B \leq 1$. We also know that $k \leq \frac{Bb^2}{4c}$. Hence the first part of the expression is positive and strictly positive for $k < \frac{Bb^2}{4c}$. As both the first and the second part of (3.11) is positive $U_i^F - U_i^P$ is positive and $U_i^F \geq U_i^P$.

Now we turn to $U_i^P \geq U_i^S$

$$U_i^P - U_i^S = \frac{16(1-B)kc + dBb^2\,(2-B)\,(4-d)}{16}\,\frac{b^2Bd^2 + 16kc - 4b^2B}{b^2Bd^2c} \tag{3.12}$$

The partial specialization equilibrium only exists for positive values of $b^2Bd^2 + 16kc - 4b^2B$, since $\frac{Bb^2}{4c} - \frac{b^2Bd^2}{16c} \leq k$ is a condition for existence. Hence, the second part of (3.12) is positive and strictly positive for $\frac{Bb^2}{4c} - \frac{b^2Bd^2}{16c} < k$. The first part of (3.12) is positive, since $B \leq 1$ and $d \leq 2$. As both the first and the second part of (3.12) is positive $U_i^P - U_i^S$ is positive and $U_i^P \geq U_i^S$.

This concludes the proof.■

Proposition 3.2. *For parameter values where the full specialization equilibrium exists, $k \leq \frac{Bb^2}{4c}$ and the return to human capital investments is positive for $l^m = 1$, $d \leq 1$, the utility is higher in the full specialization equilibrium than in the symmetric equilibrium, i.e. $U_i^F > U_i^S$.*

Proof

$$U_i^F - U_i^S = (2-B)\frac{Bb^2}{4c} - k + (8 - 4B + Bd - 2d)\frac{Bb^2d}{16c} \tag{3.13}$$

Using that $k \leq \frac{Bb^2}{4c}$ we see that (3.13) $\geq$(3.14)

$$(1-B)\frac{b^2B}{4c} + (8 - 4B + Bd - 2d)\frac{b^2Bd}{16c} \tag{3.14}$$

The first part of the expression is non-negative, since $B \leq 1$. The second part of the expression is positive, since $B \leq 1$ and $d \leq 1$. Since the difference $U_i^F - U_i^S$ is positive $U^F > U^S$.

This concludes the proof.■

4. Discussion and Conclusions

The main result of this paper is that there may exist both symmetric and asymmetric equilibria in the division of work between the sexes. Expectations on the division of labor may be self-fulfilling. The difference in wages due to different expectations on time out for child care in the non-symmetric equilibria is magnified by the differences in human capital investments. In the model the asymmetric equilibrium is welfare superior to the symmetric equilibrium. However, if factors outside the model are relevant to the welfare analysis, for example externalities for children or in changes in gender roles if men take more time out for child care, government intervention in order to coordinate on the symmetric equilibrium may be motivated.

In this model all, couples have the same preferences for the division of time out for child care. A natural avenue for further research would be to investigate the effects of heterogenous preferences among parents for the division of time out for child care. In the model presented in this paper, all couples have the same ranking over the equilibria. This would not be the case in a model with heterogenous preferences, however. For example a couple preferring the man to take more time out for child care would prefer to live in a society where men take more time out for child care, rather than living in a society where women take more time out for child care, since the wage and the return to human capital investments are higher for the sex that is expected to work more.

Appendix

This appendix first shows that the optimal wage contract, for an exogenous labor supply equal to one, is

$$w_i = \overline{MP} + B(T_i - \overline{T})$$

where

$$B = 1 - (1 + b\rho_\varepsilon)\sigma_\varepsilon^2/\sigma_T^2.$$

When that is done, it is easy to show that the labor supply and human capital investment decisions are the same for all individuals of the same sex, but might differ between the sexes. We will follow Lundberg and Startz (1983) very closely. Assume that the marginal product and the cost of acquired human capital are characterized by equations (4.1) and (3.1), and that employers observe MP imperfectly according to equation (3.2). Further, assume that the individual productivity parameter, a_i, and the error term in the test score, ε_i, are drawn from a bivariate normal distribution and that a_i and ε_i are uncorrelated. Perfectly competitive firms set wages equal to the expected marginal productivity for a given test score. Workers invest in human capital according to the wage offer schedule of the employers. We will now derive the equilibrium for wage offers and human capital investments. Write the human capital investment as a function of individual worker characteristics

$$X_i = \rho_0 + \rho_{a_i} a_i + \rho_{\varepsilon_i} \varepsilon_i.$$

We will be able to determine the parameter values in the human capital investment function in equilibrium. The firms' wage offer schedule is:

$$w_i = E(MP_i|T_i) = E(T_i - \varepsilon_i T_i).$$

The test score is normally distributed with the mean $\overline{T} = b\rho_0 + (1 + b\rho_a)\overline{a} + (1+b\rho_\varepsilon)\overline{\varepsilon}$ and the variance $\sigma_T^2 = (1+b\rho_a)^2\sigma_a^2 + (1+b\rho_\varepsilon)^2\sigma_\varepsilon^2$. The bivariate normal distribution of the test score and the test error has a correlation coefficient of $(1 + b\rho_\varepsilon)\sigma_\varepsilon/\sigma_T$. The expected value of ε_i for a test score T_i is given by

$$E(\varepsilon_i|T_i) = \overline{\varepsilon} + ((1 - B\rho_\varepsilon)\sigma_\varepsilon^2/\sigma_T^2)(T_i - \overline{T})$$

The coefficient of the test score is defined as $(1 - B)$ and the wage is given by

$$w_i = \overline{MP} + B(T_i - \overline{T}).$$

The marginal cost for acquiring human capital is c and the expected marginal wage increase is Bb, so

$$X_i = Bb/c.$$

All workers acquire the same amount of human capital. Since X is the same for all workers, ρ_a and ρ_ε must equal zero and $B = \sigma_a^2/\sigma_T^2$.

Let us now turn to the human capital investment decision when there are two sexes, then, as noted in section 3.1.2

$$X_i^m = \frac{l_i^m Bb(1 - d(1 - l_i^m))}{c} \tag{A1}$$

and

$$X_i^f = \frac{l_i^f Bb(1 - d(1 - l_i^f))}{c}. \tag{A2}$$

The wage difference between spouses is independent of ρ_a and ρ_ε, hence, all couples divide labor in the same way. Hence, according to (A1) and (A2) workers of the same sex acquires the same amount of human capital.

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Figure 1: Wages for men and women as a function of m_i

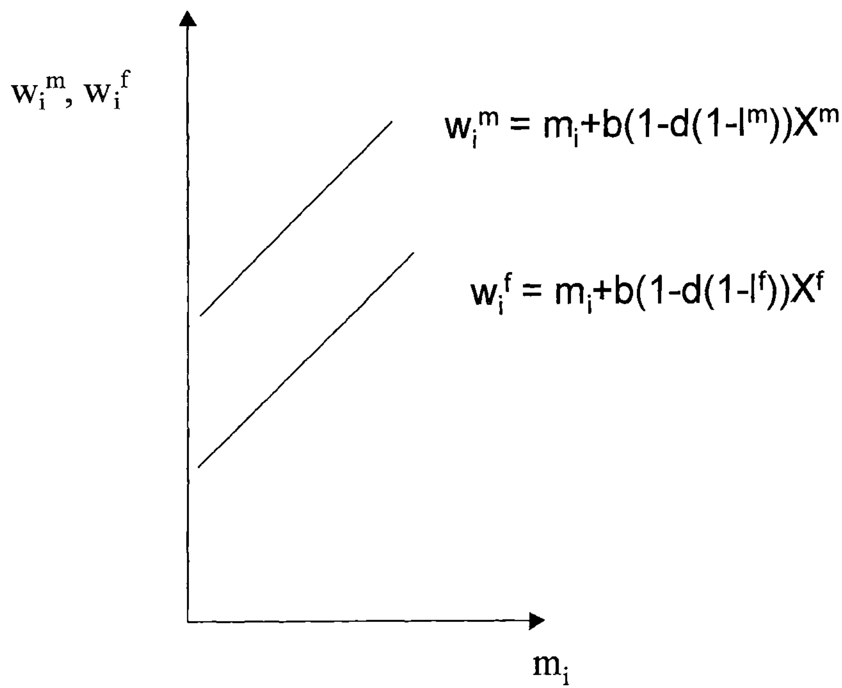
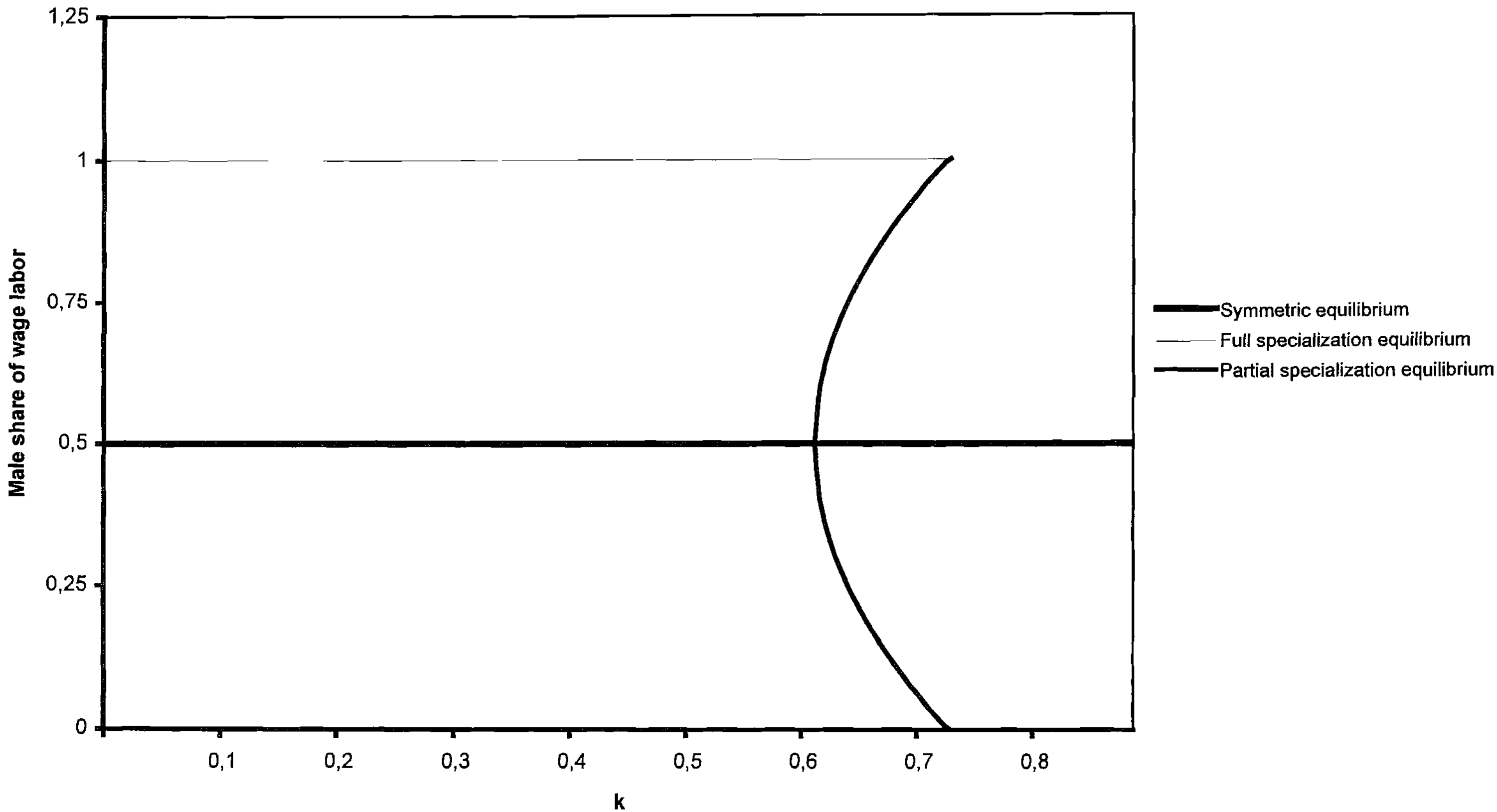


Figure 2: Male share of wage labor in different equilibria for different k



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