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MARKET SHARES, FINANCIAL CONSTRAINTS, AND PRICING BEHAVIOR IN THE EXPORT INDUSTRY

by

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INSTITUTE FOR INTERNATIONAL ECONOMIC STUDIES

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The customer market model of Phelps and Winter (1970) is used as theoretical framework for a study of Swedish exports and export prices of manufactured goods. Consistent with the model, I find that demand variations have an immediate effect on exports, while price effects take time. The export price depends on costs, exchange rates, competitors' prices, and financial conditions, measured by net borrowing. The latter effect supports the argument in Gottfries (1991) that firms set higher prices when they are financially pressed. There is evidence of short—run price rigidity in the sense that prices are set under imperfect information concerning competitors' prices and exchange rates.

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1. Introduction.

How do firms set prices in foreign markets? The answer to this question is important for the macroeconomic analysis of an open economy. It determines the consequences if the rate of inflation in one country differs from that of its trade partners, and the way in which inflation spreads from one country to another. It is also crucial for analysing the effects of a devaluation.

A closely related question is how the quantity of exports is determined. Are exporting firms able to sell any amount they want at the going "world market price" — so that exports depend mainly on supply factors in the home country — or do exporting firms face downward-sloping demand curves — so that foreign demand is an important determinant of exports? Is the relative importance of demand and supply factors different in the long run from the short run?

The purpose of this paper is to construct an empirical model of exports and export prices that can help to answer these questions. An empirical model of exports and export prices must be able to account for two well-established empirical observations. The first is that the "law of one price" does not hold for most markets — at least not in the short run — and that variations in costs are partly passed on into export prices. Thus, exporting firms have some freedom to set their own prices. A second established empirical result is that relative prices affect the quantity

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exported, but that they do so with a lag. Almost all empirical export equations contain lagged relative prices (see e.g. Goldstein–Kahn 1985, Krugman–Baldwin 1987).

The presence of lags in quantity adjustment leads one to question the theory that underlies standard equations for exports and export prices. In most empirical studies, the equations for exports and export prices are derived from a static theory, but lags are introduced when the equations are estimated. There are at least two important problems with this procedure. First, when lagged adjustment is allowed, the researcher must decide what restrictions to impose on the lag structure. Since one usually has a limited amount of data, rather tight restrictions must be imposed, but there is little theoretical basis for such restrictions. Second, if quantities respond to price changes with a lag, firms should take account of the lags when they set prices. For these reasons, it may be useful to think more explicitly about the reasons why lags appear in the equations.

A natural interpretation of slow quantity adjustment is that export markets are typically "customer markets". In such markets, each firm has a stock of customers who only gradually respond to price changes. Because of imperfect information and/or adjustment costs, customers do not immediately switch to the firm with the lowest price, but in the long run, firms with low prices gain customers, and conversely. In a customer market, the pricing decision is a dynamic optimization problem since the current

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2 Goldstein and Kahn (1985) note that estimates of polynomial lag structures are very sensitive to the restrictions imposed.

3 The original customer market model is due to Phelps and Winter (1970), and further analysis of customer markets was made by Okun (1975, 1981). The implications of consumer switching costs have been analysed by Gottfries (1986) and the literature reviewed by Klemperer (1992). Applications of this type of model to the open economy include Dohner (1984), Rodseth (1985), Gottfries (1986), and Froot and Klemperer (1989).
price affects the customer stock and future revenues.

The customer market model has interesting implications for the specification and estimation of the export equation. Furthermore, the customer market theory suggests a number of factors that may affect export prices. Interest rates, and expectations about future demand and costs may affect prices. In an earlier paper (Gottfries (1991)) I argued that when capital markets are imperfect, financial conditions may affect prices. When a firm is financially pressed, it may set a higher price so as to generate high profits now although this has negative long run effects on its market share. One purpose of this study is to examine whether variables reflecting financial conditions can help to explain prices.

Equations are estimated for Swedish exports and export prices of manufactures. The results on the quantity side support the customer market model. As expected, exports adjust slowly to prices, but quickly to variations in demand. The estimated price equation shows that costs, exchange rates, and competitors' prices are the main determinants of prices. Financial conditions, measured by net borrowing relative to value added, are found to have a large and very significant effect on the price, supporting the theory mentioned above. There is also strong evidence of short run price rigidity (predetermined prices) in the sense that firms are imperfectly informed about competitors' prices and exchange rates when they set their prices.

In Section 2 I present the data which is to be explained, and in Section 3 I outline the theory and the empirical specification of the export equation. Results for exports are presented in Section 4. Sections 5 and 6 give the specification and the results for prices. Section 7 contains a formal test of price rigidity and the final section there is some discussion of the
results.


In this section I present the basic data to be used for estimation. The data is quarterly and covers Swedish exports of manufactured goods, which includes most industrial products except food and some raw materials. (Precise definitions of the variables are given in the appendix.) Figure 1 illustrates the correlation between unit labor cost, ulc, relative to the foreign price (ulc/p*, henceforth denoted "relative cost") and the relative price of exports, p/p*. Here, p is the Swedish export price for manufactured goods and p* is a weighted index of import prices of manufactured goods for 14 OECD countries, both in Swedish currency. After a period of relatively low Swedish inflation, Swedish wages increased very much in 1975 and 1976 compared to inflation abroad, leading to an increase in relative cost. Competitiveness was restored by devaluations in April 1977 (6%) and August 1977 (10%), but then there was again some deterioration of competitiveness. Two further devaluations in September 1981 (10%) and October 1982 (16%) gave Swedish industry a big cost advantage. In the latter part of the 1980's, tight labor market conditions lead to high nominal wage increases and a steady deterioration of competitiveness, leading up to the float of the krona in November 1992.

It appears that variations in costs have a very clear impact on the relative price. Devaluations appear to have immediate effects on the relative

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4 This study builds on an empirical tradition which has been developed at the National Institute of Economic research (Konjunkturinstitutet) by Hans Olsson and others.
price, an in fact, there is a tendency of the relative price to overreact to devaluations in the short run. A natural interpretation is that export prices are partly set in kronor, and do not adjust immediately to changes in exchange rates. Similarly, the export price was relatively low in 1974 and 1977–78, when foreign prices increased rapidly, and the relatively high in 1981–82, when international prices decelerated, indicating sluggish adjustment of export prices.

Figure 2 illustrates the impact of variations in the relative price on the "market share" \((q/y)\), defined as Swedish exports to 14 OECD countries \((q)\) divided by trade-weighted imports to the same countries \((y)\).\(^5\) Both series have been smoothed by taking a five quarter moving average. In general, an increasing relative price is associated with a falling market share and conversely. There are clearly visible counter-clockwise loops. The most natural interpretation of these loops is that prices affect quantities with a lag.


What type of model of "the representative firm in the representative market" could be consistent with this data? A model with perfect competition is not consistent with the observation that variations in costs are correlated with the relative price. Of course, it may be argued that even if there is one price in each market, Swedish firms may influence the market price in the markets where they have a significant market share, so that a reduction in the aggregate relative price may reflect a change in the relative price between

\(^5\) In principle, the correct definition of the market includes also sales from domestic producers in each country, Thus, total expenditure on the relevant goods would be a better measure of market demand. Such data is more difficult to obtain, however, for the sector analysed here.
different markets. However, the magnitude of the relative price change in relation to the variation in cost suggests that this cannot be the main explanation. A model with imperfect substitutes seems more promising since it allows prices to vary between firms in the same market. A standard monopolistically competitive model does not explain sluggish quantity adjustment, however.

The customer market model of Phelps and Winter (1970) is a natural candidate to explain the above observations. In the customer market model, quantities adjust slowly because of slow diffusion of information. Each firm has a stock of customers, who only gradually learn about prices of other firms. In order to motivate my empirical equation I will use a generalization of the Phelps–Winter model where I allow goods produced by different firms to be imperfect substitutes rather than perfect substitutes as in the original Phelps–Winter model.

Assume that there is a continuum of buyers with varying preferences, who buy the good each period. Suppose that, if buyers had perfect information, the number of buyers who would purchase from the Swedish firm would be \( \alpha + \eta \frac{p_t}{p^*_t} \), where \( \eta < 0 \), and where \( p_t \) is the price charged by a Swedish firm and \( p^*_t \) is the average price in the market (both in kronor). At a particular point in time, each buyer is a customer of one particular firm. Customers know the price charged by their current suppliers, but they only gradually learn about prices of other suppliers in the market. This is captured by the following equation:

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6 The same applies to a Cournot model with perfect substitutes. See Dornbusch (1987) for a discussion of the applicability of oligopoly models to international pricing in the light of recent U.S. experience.

7 I assume that an individual Swedish firm competes with other Swedish firms to the same extent as it competes with foreign firms with the same market share. This appears to be a reasonable first approximation.
\[ x_t = (\alpha + \eta \frac{p_t}{p^*_t})^\lambda x_{t-1}^{1-\lambda}, \]

showing how the stock of customers of the Swedish firm, \( x_t \), evolves over time. Microfoundations for this type of equation have been developed by Phelps and Winter (1970) and Gottfries (1986, 1991). Now let demand per customer be \((1+\epsilon - \epsilon p_t/p^*_t) y_t^\sigma u_t\), where \( y_t \) is an observable variable (foreign imports) which shifts the demand per customer and \( u_t \) is an unobservable shock. Then demand for the firm's exports is

\[ q_t = (1+\epsilon - \epsilon p_t/p^*_t) y_t^\sigma x_t u_t. \]

The stock of customers is not observable in the available data. To get an equation that can be estimated, use (1) to substitute for \( x_t \) in (2) and then (1) dated \( t-1 \) to substitute for \( x_{t-1} \). The resulting equation for exports is

\[ \frac{q_t}{y_t^\sigma} = (1+\epsilon - \epsilon p_t/p^*_t) (\alpha + \eta \frac{p_t}{p^*_t})^\lambda \]

\[ (q_{t-1}/(1+\epsilon - \epsilon p_{t-1}/p^*_{t-1}) y_{t-1}^\sigma u_{t-1}^{1-\lambda})^{1-\lambda} u_t u_{t-1}^{\lambda-1}, \]

which can be thought of as a generalised market share equation. According to the customer market model, the relevant state variable is not the market share \((q_{t-1}/y_{t-1})\) however, but the customer stock. The customer stock differs from the market share because demand per customer varies with the price and with the unobservable shock \( u_t \), and because the demand elasticity \( (\sigma) \) may differ from unity.\(^8\) This has implications for the method of

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\(^8\) The fact that market share equations are problematic if exports from different countries have different income elasticities has been known for a long time — see
estimation which will be discussed when we turn to estimation.

The functional form is chosen so that the price elasticity is independent of demand shocks, and increasing (in absolute value) in the relative price. This functional form is chosen for two reasons. First, it is consistent with evidence (see below) that the export price is affected by variations in the foreign price, but not by variations in demand. Second, the estimated within quarter elasticity is well below unity. This is inconsistent with a demand curve that is log-linear (constant-elastic) in the price since the firm could then raise the price to infinity and make infinite profits in one quarter. In practice, a log-linear specification gives very similar parameter estimates and the data does not allow me to distinguish whether demand is linear or log-linear in the price.

Taking logs on both sides of equation (3) we get:

\[
\tilde{q}_t = \log(1+\epsilon-c p_t/p_t^s) + \sigma \tilde{y}_t + \lambda \log(\alpha + \eta p_t/p_t^s) \\
+ (1-\lambda) \{q_{t-1}-\log(1+\epsilon-c p_{t-1}/p_{t-1}^s) - \sigma y_{t-1}\} + e_t, \tag{4}
\]

where

\[
e_t = u_t - (1-\lambda) u_{t-1}, \tag{5}
\]

and where \(\sim\) above a variable denotes the log of that variable. There are

Junz and Rhomberg (1973) and Goldstein and Khan (1985). The model presented here is qualitatively similar to the models used by Krugman and Baldwin (1987) and Bean (1988).

9 For discussions of how pricing behavior depends on the form of the demand function in a static model, see Knetter (1989) and Marston (1990). Similar logic carries over to this dynamic model.
several reasons why least squares estimation of this equation would lead to biased estimates. First, there is the standard simultaneity problem: if the price depends on demand it will be positively correlated with \( u_t \). Second, measurement errors in prices are very likely because of aggregation problems etc. and prices are used as deflators in the calculation of \( q_t \) and \( y_t \). Third, since \( u_{t-1} \) is correlated with \( q_{t-1} \), a least squares estimator will produce a biased estimate of \( \lambda \) and hence of \( \eta \).

In addition, there is probably some serial correlation in \( u_t \). In fact, there is very much randomness from quarter to quarter in the export series. A likely explanation is random time allocation of exports due to weekend effects, errors in the registration of exports etc. Suppose that \( \tilde{u}_t = \epsilon_t + \xi_{t-1} - \xi_t \), where \( \xi_t \) is a shock that reallocates exports from quarter \( t \) to \( t+1 \), and that both \( \epsilon_t \) and \( \xi_t \) are i. i. d. Then \( e_t \) is a moving average of the second order:

\[
e_t = \epsilon_t - (1-\lambda) \epsilon_{t-1} + (2-\lambda) \xi_{t-1} - \xi_t - (1-\lambda) \xi_{t-2},
\]

so that \( E(e_t e_{t-1}) < 0 \), \( E(e_t e_{t-2}) > 0 \), and \( E(e_t e_{t-j}) = 0 \) for \( j > 2 \). Since \( e_t \) is negatively correlated with \( e_{t-1} \) and hence \( q_{t-1} \), a least squares estimator is likely to overestimate \( \lambda \), and hence underestimate \( \eta \) (since \( \eta \) is multiplied by \( \lambda \)). Intuitively, we may think of the term in curly brackets in (4) as an imperfect measure of the lagged customer stock \( (x_{t-1}) \), with measurement error \( u_{t-1} \). This measurement error is correlated with \( q_{t-1} \), leading to a biased estimate of \( \lambda \) if the equation is estimated by least squares. Thus, \( q_{t-1} \) should not be included in the list of instruments. Similarly, since \( q_{t-2} \) is positively correlated with \( e_t \), inclusion of \( q_{t-2} \) among the instruments may lead to a negatively biased estimate of \( \lambda \) and an overestimation of the
long-run elasticity.

For these reasons, the equation was estimated by generalized method of moments (GMM) estimation,\textsuperscript{10} taking account of second order moving average errors and conditional heteroscedasticity, using the following vector of instrumental variables: constant, seasonal dummies, trend, seasonal dummies multiplied by trend, \( u \vec{c}_t, u \vec{c}_{t-1}, \vec{p}_t^*, \vec{p}_{t-1}^*, \vec{y}_t, \vec{y}_{t-1}, \vec{y}_{t-3}, \vec{q}_{t-3}, \) strike dummy and lagged strike dummy.\textsuperscript{11} In the following, this vector is denoted \( Z_t^q \). Precise definitions of the variables are given in the Appendix. All indexes were normalised to unity in the first quarter of 1980 so that \( \epsilon \) and \( \eta \) are price elasticities at that point.

4. Exports: Results.

Line 1 in Table 1 shows the parameter estimates with \( Z_t^q \) as instrument vector. The autocorrelations for the residual are listed in the note to the table. Note that only the first two autocorrelations are significant, and that the signs are those one would expect. The estimate of \( \lambda \) indicates slow adjustment of the customer stock: about 60 percent of the adjustment occurs within one year and about 80 percent of the adjustment has happened after two years.

Conventional specification and estimation often leads to very low

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\textsuperscript{10} In this context, the GMM estimator is best thought of as a generalisation of two stage least squares (2SLS) that takes account of moving average errors as well as heteroscedasticity conditional on the instruments. Under these conditions, 2SLS would produce consistent estimates of the parameters, but the standard errors would be biassed. In practice, the results of 2SLS and GMM are very similar.

\textsuperscript{11} I used unit labor cost, rather than the full cost index, \( w_t \), as instrument since \( w_t \) contains Swedish producer prices and Swedish import prices, both of which are likely to be simultaneously determined.
long—run price elasticities (see Goldstein and Kahn (1985)). The theory outlined above suggests that this may be due to incorrect dynamic specification and estimation. The choice of instruments matters very much for the estimated long run elasticity. If the equation is estimated by nonlinear least squares or if $q_{t-3}$ is replaced by $q_{t-1}$ in the instrument list, the long—run elasticity, $\eta$, is estimated to be close to unity (lines 2 and 3). If, on the other hand, $q_{t-3}$ is replaced by $q_{t-2}$ in the instrument list, the estimated long run elasticity is very large (line 4). These results are consistent with the theoretical prediction that $\lambda$ will be overestimated ($\eta$ underestimated) in lines 2 and 3 and underestimated ($\eta$ overestimated) in line 4. Note, however, that the within—quarter effect ($\epsilon + \lambda \eta$) is around one half in all cases, so the difference concerns primarily the long run effect.

It may also be of interest to compare the model to a more general dynamic specification. In order to do this, consider, for convenience, a log—linear version of the model:

$$\tilde{q}_t = (\epsilon + \lambda \eta)(\tilde{p}_t - \tilde{p}^*_t) + \sigma \tilde{y}_t + (1-\lambda) \{q_{t-1} - \epsilon (\tilde{p}_{t-1} - \tilde{p}^*_{t-1}) - \sigma \tilde{y}_{t-1}\} + \epsilon_t.$$  

(7)

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12 See Goldstein and Kahn (1985) for a review of the empirical estimation of export and import functions. Lundborg (1981), following Goldstein and Kahn (1978), found a price elasticity of $-48$ for Swedish exports. One reason for the difference is probably that Lundborg used data for total Swedish exports, leading to more severe aggregation problems.

13 These results support the view that the "measurement error" with respect to the lagged customer stock ($u_{t-1}$) matters. As we will see below, the export price appears largely independent of demand factors, so the textbook simultaneity problem appears less important in practice.

14 Bean (1988) estimated a similar equation for British exports (annual data). He could not reject the hypothesis that the long run price elasticity is infinite, so that a period with high relative price leads to a permanent loss of market share.
According to this model, we should have a long run equilibrium relationship:

$$\tilde{q}_t = (\epsilon + \eta)(\tilde{p}_t - \tilde{p}_t^*) + \sigma \tilde{y}_t.$$  \hfill (8)

We can then write an "error correction" equation for exports:

$$\Delta \tilde{q}_t = -\lambda (\tilde{q}_{t-1} - \beta (\tilde{p}_{t-1} - \tilde{p}_{t-1}^*) - \sigma \tilde{y}_{t-1}) + a_1 (\Delta \tilde{p}_t - \Delta \tilde{p}_t^*)$$

$$+ a_2 (\Delta \tilde{p}_{t-1} - \Delta \tilde{p}_{t-1}^*) + a_3 \Delta \tilde{y}_t + a_4 \Delta \tilde{y}_{t-1} + a_5 \Delta \tilde{q}_{t-1} + \epsilon_t. \hfill (9)$$

If the customer model is correct we would expect the following coefficient estimates for the error-correction equation:

$$\lambda \cong .21 \quad \beta = \epsilon + \eta \cong -1.6 \quad \sigma \cong .85 \quad a_1 = \epsilon + \lambda \eta \cong -.44$$

$$a_2 \cong 0 \quad a_3 = \sigma \cong .85 \quad a_4 \cong 0 \quad a_5 \cong 0. \hfill (10)$$

Estimation of this equation by GMM, adding $\tilde{u}_t c_{t-2}$, $\tilde{p}_{t-2}^*$ and $\tilde{y}_{t-2}$ to the list of instruments, $Z_t^q$, to improve identification, gives the following estimates:

$$\lambda = .168 \quad \beta = -1.701 \quad \sigma = .862 \quad a_1 = -.469$$

$$\quad (.096) \quad \quad (.167) \quad \quad (.039) \quad \quad (.156)$$

$$a_2 = -.029 \quad a_3 = .824 \quad a_4 = -.095 \quad a_5 = -.091$$

$$\quad (.222) \quad \quad (.122) \quad \quad (.237) \quad \quad (.242) \hfill (11)$$
All the parameters values are close to those predicted by the customer market model. In particular, the estimate of $a_3$ supports the prediction of the customer market model that exports respond immediately to variations in demand.\footnote{A traditional Koyck specification, for example, would impose the same speed of adjustment with respect to prices and demand.}

So far, we have discussed these estimates without consideration of the trends in the variables and their implications for the estimates reported above. Augmented Dickey–Fuller (ADF) tests (including trend) of the hypotheses that $\ddot{q}$ and $\ddot{y}$ are integrated of order one (I1) are $-2.30$ for $\ddot{q}$ and $-3.27$ for $\ddot{y}$. The first value is not significant and the second is significant on the ten percent level (Davidson–MacKinnon, 1993). Thus, we accept the hypothesis that exports are (I1), but foreign imports appear to be trend–stationary.

If the long–run price elasticity is infinite, the relative price ($\ddot{p}-\ddot{p}^*$) should be stationary. An ADF test (excluding trend) of the hypothesis that the relative price is not stationary gave a test statistic of $-2.66$, which is just significant on the 10 percent level. One may argue that the latter part of the period was a clear disequilibrium situation, however. If the last years are excluded, and the test is run for the period 1971:1 to 1987:4, the ADF statistic becomes $-3.0$, which is significant on the 5 percent level, indicating that the relative price may in fact be stationary.

What are the implications of the trends for the parameter estimates reported above? Note that if (8) is really a cointegrating relationship, then all the parameter estimates have standard distributions. This follows from the fact that if (8) is a cointegrating relationship, all the coefficients can be
written as coefficients for stationary variables (Sims, Stock and Watson (1990)). To see whether there is evidence of cointegration, I generated the residual from (8), using the parameter estimates in (11):

\[ \text{res}_t = q_t + 1.70 (p_t - p_t^*) - 0.86 \gamma y_t, \]  

and calculated the ADF test (without trend) for this variable. The test statistic was -2.62, which is far from significant. Thus, we cannot reject the hypothesis that \( \text{res}_t \) is non-stationary, so that (8) is not a cointegrating relationship. This may be due to low power of the test, however.\(^{16}\)

What can we conclude from the examination of the trends? One possibility is that there is a downward sloping long-run demand function, the relative price is II, and the reason why we do not reject non-cointegration is weak power of the ADF test.

On the other hand one may have an a priori view that the long-run elasticity should be very large since Swedish exports typically have very close substitutes abroad. This view is supported by the above mentioned test statistics, indicating that the relative price may be stationary. If this view is correct, (8) is in fact not a cointegrating relationship. One may then estimate a model where an infinite long-run price elasticity is imposed. If we replace (1) by

\[ x_t = (\phi + \gamma p_t/p_t^*) x_{t-1} \]  

\(^{16}\) Note that in this test, the null hypothesis is that there is no cointegration. If this hypothesis is not rejected, it does not mean that cointegration has been rejected.
we get an export equation of the form

\[ \tilde{q}_t = \log(1+\epsilon \frac{e p_t}{p_t^*}) + \sigma \tilde{y}_t + \log(\phi + \gamma \frac{p_t}{p_t^*}) \]

\[ + \{q_{t-1} - \log(1+\epsilon \frac{e p_{t-1}}{p_{t-1}^*}) - \sigma \tilde{y}_{t-1}\} + \epsilon_t. \] (14)

Estimation of this equation, using $Z_t^Q$ as instruments, gives the estimates reported on line 5 in Table 1. Note that the short-run implications of this equation are quite similar to those obtained above, but the long-run implications are very different.


Now consider pricing behavior. Because the price affects the customer stock, and hence future revenues, the pricing decision is a dynamic optimization problem. Assume that marginal cost is $\beta w_t$, where $w_t$ is a composite price of factors of production. Then the objective of the firm is to set the price so as to maximize:

\[ E_t \sum_{j=0}^{\infty} \delta^j (p_{t+j} - \beta w_{t+j}) (1+\epsilon \frac{e p_{t+j}}{p_{t+j}^*}) y_{t+j}^\sigma x_{t+j} u_{t+j} \] (15)

subject to (1) or (13).17 I am not able to derive a closed for solution to this optimization problem, but in an earlier paper (Gottfries 1986) I derived an

17 When the marginal cost is independent of the quantity produced, pricing in the export market can be analysed independent of sales in the domestic market and inventory accumulation.
approximate solution for a similar model and used the approximate solution to find out how various variables affect the price.

Increases in costs and market prices should raise the price in the same period. Prices should also depend on expectations about the future. Higher expected future costs should raise the price ("normal cost pricing"). An expected future increase in the market price should lead to a lower price today since customers are more valuable if they pay more in the future. If, as we assumed above, the marginal cost curve is flat, a permanent increase in demand should not affect the price. A temporary increase in demand should raise the price, and an expected future increase in demand should reduce the price (again because customers become more valuable). An increase in the interest rate should have a positive effect on the price since the firm sets a higher price if it discounts the future more.18

If credit markets are imperfect, financial variables may also affect the price, as suggested by Gottfries (1991). If a firm is financially pressed, it may set a higher price in order to increase profits now although this has negative consequences for future profits. I use net borrowing relative to value added, $nb_t$, to measure financial conditions. The idea is that if the firm has to borrow a lot of new money every month, it is more anxious to raise profits now at the expense of future profits. High net borrowing may reflect low operating profits, a large stock of financial debt, or high investment expenditure which needs to be financed. In practice, net borrowing is highly correlated with investment expenditure.

One alternative would be to estimate a tightly specified structural price equation which is derived from the intertemporal choice problem of the

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18 The role of the interest rate for the markup is emphasised by Fitoussi and Phelps (1988) and Phelps (1994).
firm. I did not follow this approach because only an approximate solution can be derived, and because the estimates may be sensitive to misspecification of the processes driving the exogenous variables etc.\(^\text{19}\) Instead, the theory was used more loosely to suggest which variables should be included in the price equation.

Experiments were made with variables measuring demand, future costs, future market prices, capacity utilization and interest rates, but these coefficients were found to be insignificant and/or unstable.\(^\text{20}\) I therefore estimated a simple price equation with costs, \(w_t\), exchange rate, \(e_t\), foreign import price in foreign currency, \(p_t^f\) and net borrowing, \(nb_t\), as explanatory variables.

It is quite likely that firms are imperfectly about competitors' prices and other variables when they set prices for a particular period. Hence, prices depend on firms' expectations about these variables. Under rational expectations, one can estimate the price equation with actual values of the variables, however, provided that one uses instruments which are known to firms when they set their prices. Allowing for prices to be set up to a year in advance, I used the following instrument vector: constant, seasonal dummies, trend, \(\tilde{u}_t\), \(\tilde{c}_{t-4}\), \(\tilde{c}_{t-6}\), \(e_{t-4}\), \(p_{t-4}^f\), \(p_{t-6}^f\), \(nb_{t-4}\), \(p_{t-4}^o\), where \(p_t^o\) is the Swedish import price for oil. In the following, this vector is denoted

\(^{19}\) Blanchard and Melino (1986) estimated a very tightly specified structural price equation which was derived from the intertemporal choice problem of the firm. Another alternative is to estimate an Euler equation for the pricing problem of the firm. I have earlier made attempts in this direction (Gottfries 1985, 1988). Unfortunately, the Euler equation for this problem is quite complicated and hence the results are hard to interprete.

\(^{20}\) The insignificance of forward-looking variables is negative for the customer market model. One possibility is that firms have little knowledge about future changes in costs and market prices. It is not obvious how to define the real interest in this context, and the coefficient for the interest rate appeared to be quite sensitive to the specification.
$Z_{t-4}$.

6. Prices: Results.

If prices are set in advance, the price equation will have a moving average error. Therefore, the equation was estimated by GMM, allowing for fourth order moving average errors and conditional heteroscedasticity. The result was as follows:

$$
\tilde{p}_t = 0.368 \tilde{w}_t + 0.689 \tilde{e}_t + 0.623 \tilde{p}_t^f + 0.341 \text{nb}_t \quad (16)
$$

s. e.: .0214 $\quad R^2 : .998.$

Constant and seasonals were included in the equation but they are not reported. The estimation period was 1971:1–1990:4. Theory suggests that the coefficients for the foreign price and the exchange rate should be the same and that this coefficient should be one minus the coefficient on the cost index (nominal neutrality). This indeed appears to be the case, and imposing these constraints we get the following equation:

$$
\tilde{p}_t = 0.406 \tilde{w}_t + 0.594 (\tilde{e}_t + \tilde{p}_t^f) + 0.308 \text{nb}_t \quad (17)
$$

s. e.: .0207.

As expected, there is considerable positive serial correlation. The autocorrelations for the residual are: .507 (.112), .121 (.138), .011 (.139), .097 (.139), .019 (.140). Note that only the first autocorrelation is significant,
consistent with the view that pre-set prices cause a moving average error.

The equation explains about 90 percent of the variation in the relative price. The effect of net borrowing (nb) is not only strongly significant, but it is also quantitatively important. Over the course of the business cycle net borrowing may vary from, say, zero to .15, implying a change in price of about five percent.

7. A Test of Price Rigidity.

So far, I have allowed for the possibility that prices are pre-set (and/or firms have lagging information) by using lagged instruments for estimation. In this section, an explicit test of price rigidity is made, based on ideas in Gottfries-Persson (1988) and Gottfries-Persson-Palmer (1989). Price rigidity is taken to mean that prices do not reflect the most recent information. Consider, for illustrative purposes, a simplified price equation:

\[ p_t = b_0 + b_w E_t(w_t) + b_p E_t(p_t^*) + \mu_t, \]  

(18)

where \( E_t \) denotes the expectation conditional on the information that firms have when they set prices for period \( t \) and \( \mu_t \) is a stochastic term reflecting other factors which affect the price. Let us now decompose movements in costs, \( w_t \), into those predictable on the basis of lagged information, and those which could not be predicted. Assume that firms know a vector of lagged variables, \( Z_{t-j} \) and let \( Q_t \) be a vector of variables which they may know when they set prices for period \( t \). \( Q_t \) may include market price and cost in

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21 Giovannini (1988) and Marston (1989) also distinguish planned and unplanned variations in prices. The methodology proposed here can be regarded as a generalization of the methodology used by Marston (1989).
period \( t \) etc. Projecting \( E_t(w_t) \) recursively on \( Z_{t-j} \) and the projection of \( w_t \) on \( Z_{t-j} \) and \( Q_t \) we get\(^{22}\)

\[
P(E_t(w_t)|Z_{t-j}, P(w_t|Z_{t-j}, Q_t))
\]

\[
= P(E_t(w_t)|Z_{t-j}) + m_w[P(w_t|Z_{t-j}, Q_t) - P(P(w_t|Z_{t-j}, Q_t)|Z_{t-j})]
\]

\[
= P(w_t|Z_{t-j}) + m_w[P(w_t|Z_{t-j}, Q_t) - P(w_t|Z_{t-j})].
\] (19)

The first equality is the "law of iterated projections" (See Sargent 1979) and the second equality follows from the assumption that \( Z_{t-j} \) is known by agents. We can think of \( m_w \) as a measure of the "information advantage" that firms have concerning \( w_t \), relative to the information embodied in \( Z_{t-j} \). If agents know \( Q_t \) when they set their prices for period \( t \), \( m_w \) is unity. If firms are imperfectly informed about \( Q_t \) when they set their prices \( m_w \) is lower than unity. If agents have no information about \( w_t \) beyond \( Z_{t-j} \), \( m_w \) is zero.

Consider now how \( m_w \) can be estimated. Write \( P(w_t|Z_{t-j}) \) as \( Z_{t-j} \alpha_w \), where \( \alpha_w \) is a vector of coefficients, and let

\[
\zeta_{wt} \equiv E_t(w_t) - P(E_t(w_t)|Z_{t-j}, P(w_t|Z_{t-j}, Q_t)),
\] (20)

and

\[
e_{wt} \equiv w_t - P(w_t|Z_{t-j}, Q_t).
\] (21)

\(^{22}\) I assume that all variables are normally distributed so that linear projections equal conditional expectations.
We then have

$$E_t(w_t) = Z_{t-j} \alpha_p + m_w[w_t - Z_{t-j} \alpha_w - e_{wt}] + \zeta_{wt}. \quad (22)$$

Now use this equation to substitute for $E_t(w_t)$ in the price equation. Similarly, projecting $E_t(p^*_t)$ on $Z_{t-j}$ and $P(p^*_t | Z_{t-j}Q_t)$, and using the result to substitute into the price equation, we get the following system of equations:

$$p_t = b_0 + b_w[Z_{t-j} \alpha_w + m_w(w_t - Z_{t-j} \alpha_w)] + b_p[Z_{t-j} \alpha_p + m_p(p^*_t - Z_{t-j} \alpha_p)]$$

$$- b_w m_e e_{wt} + b_w \xi_{wt} - b_m p^* e_{pt} + b_p \xi_{pt} + \mu_t. \quad (23)$$

$$w_t = Z_{t-j} \alpha_w + \epsilon_{wt} \quad (24)$$

$$p^*_t = Z_{t-j} \alpha_p + \epsilon_{pt}. \quad (25)$$

By definition

$$E[\xi_{wt} | Z_{t-j}, P(w_t | Z_{t-j}Q_t)] = E[\xi_{pt} | Z_{t-j}, P(p^*_t | Z_{t-j}Q_t)] = 0 \quad (26)$$

and

$$E(e_{wt} | Z_{t-j}Q_t) = E(e_{pt} | Z_{t-j}Q_t) = E(\epsilon_{wt} | Z_{t-j}) = E(e_{pt} | Z_{t-j}) = 0 \quad (27)$$

and I assume that
\[ E[\xi_{wt} | Z_{t-j}^t, P(w_t | Z_{t-j}^t, Q_t), P(p_t^* | Z_{t-j}^t, Q_t)] = \]

\[ E[\xi_{pt} | Z_{t-j}^t, P(w_t | Z_{t-j}^t, Q_t), P(p_t^* | Z_{t-j}^t, Q_t)] = 0. \] (28)

This assumption is more reasonable the more unrelated \( w_t \) and \( p_t^* \) are in the short run. A sufficient but not necessary condition is that the innovations (relative to \( Z_t \)) in \( w \) and \( p^* \) are independent.

If we have chosen \( Z_{t-j} \) and \( Q_t \) so that \( E(\mu_t | Z_{t-j}^t, Q_t) \), the system of equations (24)–(26) can be estimated by three stage least squares, using \( Z_{t-j} \) and \( Q_t \) as instrumental variables.\(^{23}\) If we find, for example, that \( m_w \) is significantly lower than unity, this is evidence of price rigidity: prices do not reflect the most recent information about \( w_t \).

This approach was used to estimate information advantage coefficients with respect to the explanatory variables in the price equation. For \( Z_{t-j} \) I used \( Z_{t-4} \), the vector of instrumental variables mentioned above, and \( Q_t \) was specified as the four quarter change in (log of) hourly wage cost, dated \( t \), \( p_t^* \), \( e_t \) and \( n_b_t \). The resulting equation turned out as follows:

\[
\bar{p}_t = .353 \left[ Z_{t-4} \alpha_w + 1.573 (w_t - Z_{t-j} \alpha_w) \right]^{(441)} \\
+ .647 \left[ Z_{t-4} \alpha_p + .431 (p_t^* - Z_{t-j} \alpha_p) \right]^{(180)} \\
+ .647 \left[ Z_{t-4} \alpha_v + .254 (v_t - Z_{t-j} \alpha_v) \right]^{(93)}
\]

\(^{23}\) Note that while \( \xi_{wt} \) is not orthogonal to \( Z_t \) and \( Q_t \), it is orthogonal to the instrument for \( w_t \) by definition.
\[ + .353 \left( Z_{t-4} \alpha_{nb} + .013 (nb_t - Z_{t-1} \alpha_{nb}) \right) \]

\begin{align*}
&\text{s. e.: .0155.}
\end{align*}

The projection equations \((w_t = Z_{t-4} \alpha_w \text{ etc.})\) are not presented since they are of little interest. The estimated information coefficients indicate that firms have imperfect information concerning the market price, the exchange rate and the financial situation when they set prices. In this sense, there is significant price rigidity. The information coefficient with respect to the cost variable is larger than (but not significantly different from) unity. Firms appear to be better informed about their own costs than about competitors' prices and exchange rates when they set their prices.

These tests indicate that firms have imperfect information about conditions in period \(t\) when they set prices for that period. Formally, we cannot say whether this is because prices are set at an earlier point in time, or firms have lagging information. The fact that the information coefficient for the exchange rate is rather low supports the pre-set-prices interpretation, however, since information concerning exchange rates is available every day in the newspaper. A natural interpretation of the results is that prices are set a few quarters in advance, firms have advance knowledge of costs, and exchange rate changes are relatively unpredictable.\(^{24}\)

The results imply that export prices in kronor do not adjust immediately to exchange rate changes. Hence, the pass-through of exchange

\(^{24}\) This interpretation is consistent with the results of a questionnaire study by Assarsson (1989) about the pricing practice of Swedish industrial firms. Assarsson found that one third of the companies changed prices at most once a year, another third changed prices up to twice a year.
rate changes to export prices in foreign currency is larger in the short run than in the long run (c. f. Figure 1). This result is similar to what Knetter and Gagnon (1990) found for most of the export prices they considered, but opposite to Hooper and Mann (1989) who found short-run pass-through to be lower than long-run pass through.25

8. Discussion.

The results on the quantity side suggest that the dynamic aspects of demand are an important part of the firm's environment. This is important not only for international economics, but also for general macroeconomic and microeconomic theory. If customers react slowly to price changes in international markets, they probably do so in the market for haircuts. Of course, any business magazine offers plenty of casual evidence that firms are aware of the dynamic aspects of demand and that they view the market share as an important asset of the firm.

A long-standing issue in industrial organization is whether firms compete in prices or quantities. In the theoretical analysis above, the Swedish firm is assumed to be small relative to the market, so the market price does not depend on what the firm does. Hence, there is no strategic difference between price and quantity competition. If competitor's prices and demand were were known to the firm, choosing price would be equivalent to choosing the quantity exported. But if the firm is imperfectly informed about conditions in period t when it takes its decision for that period, setting the price is not equivalent to setting the quantity. In the analysis above, the

25 A possible reason for the difference is that exporters usually set prices (and invoice) in their own currency, except when they export to the U. S. (see Knetter (1989).
control variable of the firm was the price, and evidence was found that prices are set under imperfect information about competitors' prices. This supports the view that, in the short run, firms set prices rather than quantities. The mode of competition is Bertrand, and demand is very inelastic in the short run.

The estimated price equation shows that variations in costs are partly passed on into prices. The pass-through of exchange rate changes and variations in costs is about 35–40 percent. Firms appear to be better informed about costs than about competitors' prices, so cost have a more immediate effect on prices. There is strong evidence that firms are imperfectly informed about competitors' prices, supporting theories emphasizing pre-set prices under imperfect information, e.g. Gordon (1981), Andersen (1985) and Nishimura (1986).

The financial situation of the firm is important for price setting. This finding may help towards an understanding of the cyclical properties of prices. Net borrowing tends to lag behind capacity utilisation by about four quarters and it is highly correlated with fluctuations in investment. A natural interpretation is that firms decide to invest when capacity utilisation is high, and investment then continues for a while because it takes "time to build." According to the estimates above, net borrowing affects prices with about a four quarter lag. The result is an eight quarter lag between an increase in capacity utilization and the effect on the price. The result is something that may looks like very slow price adjustment, or even a countercyclical pattern of the markup. (See Bils (1987) and Rotemberg and

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26 Bhaskar, Machin and Reid (1993) and Chevalier and Scharfstein (1994) also report evidence to this effect.

27 A similar pattern is noted by Brealey and Myers (1986, p. 292).
Woodford (1991) for evidence of countercyclical markups in the U.S.)
Data appendix.

The definition of manufactured goods is SITC 5–9 except 68 (nonferrous metals) and 793 (ships) in the trade statistics, and SNR (SNI) 32, 33 excl.33111, 34 excl. 34111, 35 excl. 353 and 354, 36, 371, 38 excl.3841 and 39 in the production statistics. Thus the main exclusions from total industrial production (SNR 3) are food, sawnwood, pulp, petroleum products, nonferrous metals and ships. The 14 countries referred to below are Canada, USA, Japan, Belgium, Netherlands, France, United Kingdom, Germany, Switzerland, Austria, Italy, Norway, Denmark and Finland. KITS is the database available at the National Institute for Economic research (Konjunkturinstitutet, KI).

q: value of exports of manufactured goods to 14 countries (Source: Swedish foreign trade statistics and KITS), deflated by p.

p: "hybride" price index for exports of manufactured goods, based on a mixture of unit value and producer price indecies. Producer price indecies are based on transaction prices for a sample of goods. (Source: KITS.)

p*: weighted average of import price indecies for manufactures for 14 countries, measured in Swedish currency, constructed from OECD trade statistics, series A and national statistics. The weights are based on the shares of Swedish exports going to each country. (Source: KITS.)

y: correspondingly weighted average of imports volume indecies for manufactures for 14 countries. Volumes are calculated by deflating import values by import prices (values from OECD trade statistics, series A, and national statistics). (Source: KITS.)

e: export–weighted exchange rate index. (Source: KITS.)

\[ p^f = p^* / e. \]

ulc: unit labor cost; wage cost per hour divided by production per
hour. Wage cost per hour is .7 times wage cost per hour for blue-collar workers plus .3 times wage cost per hour for white collar workers. The latter was obtained by dividing the wage cost per month by stipulated hours per month for white collar workers. Production per hour is production divided by working hours for SNR 3. (Source for hours: National Accounts.)

\[ w: \text{cost index, constructed from ulc and an input price index constructed from producer prices, domestic deliveries, and producer prices for imports for various SNR groups. Weights are based on input-output statistics for 1980.} \]

\[ nb: \text{change in net debt in a year divided by value added for the same year, SNI 3. Net debt is short and long term debt plus .15 times untaxed reserves minus financial assets. (.15 is a rough estimate of the effective corporate tax rate.) Financial assets is total assets minus inventories, machines, real estate, imaginary assets and shares. Note that this is a yearly series. (Source: Företagen, Statistics Sweden.)} \]

the strike dummy takes the value unity in the second quarter of 1980 and zero otherwise.

capacity utilization is production relative to capital stock for SNR 3, where the capital stock is calculated as in Hansson (1991).

*Note:* \( p^* \) and \( y \) for the first half of the 1970's were calculated at KI, but are no longer available through KITS.

Seasonally unadjusted data was used for estimation and seasonal dummies were included in the equations. Figure 1 shows seasonally adjusted data, however and Figure 2 shows a five quarter moving average with interpolated value for the strike quarter 1980:2.
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Marston, R. C., 1990, Pricing to market in Japanese manufacturing, Journal of


Table 1. Results for Export Equation.

<table>
<thead>
<tr>
<th>Line</th>
<th>$\epsilon$</th>
<th>$\sigma$</th>
<th>$\lambda$</th>
<th>$\eta$</th>
<th>$\lambda\eta (\gamma)$</th>
<th>s. e.</th>
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<tr>
<td>1.</td>
<td>-.146</td>
<td>.853</td>
<td>.205</td>
<td>-1.467</td>
<td>-.300</td>
<td>.0289</td>
</tr>
<tr>
<td></td>
<td>(.167)</td>
<td>(.029)</td>
<td>(.067)</td>
<td>(.365)</td>
<td>(.075)</td>
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<tr>
<td>2.</td>
<td>-.081</td>
<td>.847</td>
<td>.315</td>
<td>-1.177</td>
<td>.371</td>
<td>.0283</td>
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<td></td>
<td>(.275)</td>
<td>(.027)</td>
<td>(.077)</td>
<td>(.341)</td>
<td>(.147)</td>
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<tr>
<td>3.</td>
<td>-.154</td>
<td>.856</td>
<td>.306</td>
<td>-1.020</td>
<td>-.313</td>
<td>.0285</td>
</tr>
<tr>
<td></td>
<td>(.242)</td>
<td>(.022)</td>
<td>(.080)</td>
<td>(.240)</td>
<td>(.132)</td>
<td></td>
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<tr>
<td>4.</td>
<td>-.525</td>
<td>.767</td>
<td>.008</td>
<td>-14.5</td>
<td>-.113</td>
<td>.0332</td>
</tr>
<tr>
<td></td>
<td>(.119)</td>
<td>(.084)</td>
<td>(.007)</td>
<td>(7.0)</td>
<td>(.056)</td>
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</tr>
<tr>
<td>5.</td>
<td>-.376</td>
<td>.729</td>
<td>-</td>
<td>-</td>
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<td>(.111)</td>
<td>(.077)</td>
<td></td>
<td></td>
<td>(.045)</td>
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Notes:
A constant and seasonals were included in all equations. Seasonal dummies multiplied by trend were included in the export equation since there are indications of a trend in the seasonals. A strike dummy was included for the second quarter 1980. The strike dummy was entered as a demand shift variable analogous to $y$. Estimation period: 1971:1–1992:3. Program: TSP 4.2.

*Line 1 shows estimates with $Z^q$ as instrument vector. Autocorrelations for residual in this estimation are $-.459 (.108), .277 (.129), -.184 (.135), .132 (.138), -.129 (.140)$.*

*Line 2 reports estimates by nonlinear least squares.*

In line 3 the instruments are $Z^q$, except that $q_{t-3}$ is replaced by $q_{t-1}$.

In line 4 the instruments are $Z^q$, except that $q_{t-3}$ is replaced by $q_{t-2}$.

*Line 5 shows results for equation (14) with $Z^q$ as instruments.*
Fig. 2 Relative price and market share

Note: The numbers mark the first quarter in each year. The series are smoothed by taking a five quarter moving average.