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# An Econometric Analysis of Exports of Metals: Product Differentiation and Limited Output Capacity

**Abstract:**
A framework based on product differentiation and limited output capacity is established as a foundation for modelling Norwegian exports of primary metals: Producers of metals are exposed to economic cycles, which lead to changes in the capacity utilisation. When output capacity is available, the price is set by the supply side and export is then given from the demand side. In periods with full capacity utilisation, export is constrained from the supply side and we are not able to identify the demand curve. Using data on capacity utilisation, supply and demand variables in an export-equation are weighted in accordance to the development in economic cycles. We estimate long-run solutions including foreign demand, relative prices and capital stock.

**Keywords:** capacity constraint, export modelling, metals, product differentiation

**JEL classification:** F14

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# 1. Introduction

The most common approach in empirical analyses of the export determination is to estimate a demand function including foreign income and relative prices as exogenous variables, assuming that goods produced in different countries are differentiated [cf. Armington (1969)]. Sometimes a supply relation is estimated simultaneously with, or as an alternative to, the demand function, as for instance in Lindquist (1993), Muscatelli et al. (1995), Naug (1995a/b) and Athukorala and Riedel (1996). Incorporating the supply side of the economy into the model may be important both to avoid an omitted-variable bias and to disentangle the influence of supply side factors from the price elasticity estimates [see e.g. Naug (1995b) and Madsen (1998)]. A comprehensive survey of methodological issues and results from empirical trade studies are contained in Goldstein and Kahn (1985).

The sizes of income and price elasticities have been subject to a great deal of attention in the empirical macroeconomic literature. Goldstein and Kahn (1985) do not find empirical evidence that there have been major changes in the size of estimated elasticities of income and relative prices over different time periods. This parameter stability supports the use of a standard demand model when describing long-run trade flows. Harberger (1957) presents a consensus export price elasticity of "...near or above -2". Goldstein and Kahn suggest a consensus estimate within a range of -1.25 to -2 for exports in general. They also find that there are significant differences in both price and income elasticities across commodity groups, and that the price elasticity for manufactures is significantly larger than for non-manufactures.

Econometric studies by Lindquist (1993) and Naug (1995a) show that both demand and supply variables contribute to explain the development of Norwegian exports of primary metals. Lindquist estimates demand and supply models separately, and finds it difficult to discriminate between them on statistical ground. She suggests a modelling strategy where demand and supply variables are included in the same model. Naug estimates a mixed model, and finds support for this approach. Norwegian exports of metals are dominated by relatively homogenous products. According to economic theory, this should be reflected in high price elasticities of demand. However, in previous studies of Norwegian exports of metals, the estimated price elasticities are surprisingly low.

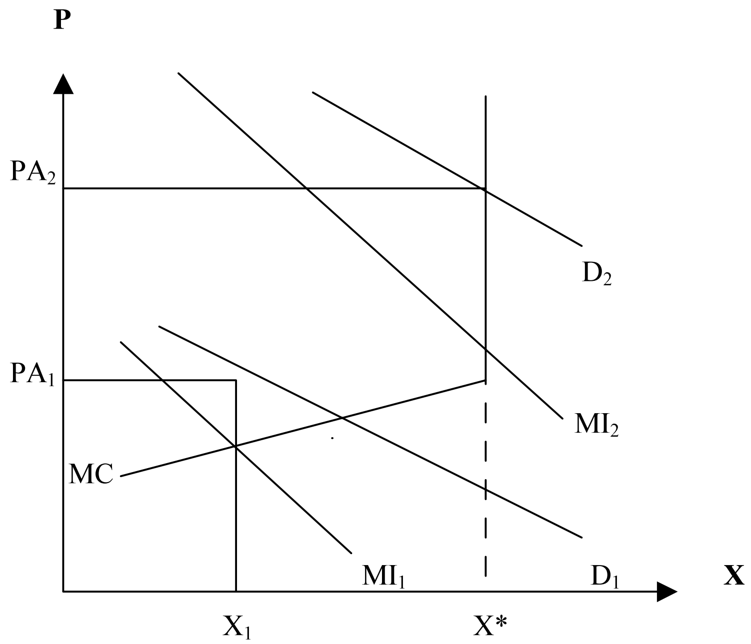
In this paper, we assume that the world market is characterised by imperfect competition. The market consists of firms producing differentiated products. Each producer faces a downward-sloping demand curve and acts as a monopolist. We put forward the hypothesis that the low estimates of price elasticities for Norwegian exports of metals are due to an identification problem. Within a conventional export demand framework, we are not able to estimate the “true” demand curve because export is not determined by demand (given the price) at all times. As long as the output capacity is not fully utilised, the price is set by the supply side and export is determined by demand. Exporters of metals are, however, facing extensive fluctuations in demand. These are capital-intensive industries, and expansion in the output capacity is costly and takes time. This makes it practically impossible to make the output capacity follow all changes in demand. As capacity utilisation rises, a switching point will be reached where export is constrained by output capacity and the demand variables "lose" their explanatory power. In other words, the export determination switches from the demand side to the supply side when output capacity is fully utilised. In order to identify the “true” demand curve, and thereby the correct price elasticity, we have to identify the periods where exports actually are determined by demand. To overcome this identification problem, we include demand and supply variables in the same model for Norwegian exports of metals, weighted in accordance to the activity level in the market. A measure of capacity utilisation in the metal industry is used to construct the weights.

We estimate two versions of the model, using discrete and continuous weights respectively. The two versions represent two alternative hypotheses. The use of discrete weights is consistent with the assumption that all firms face the same regime at all times and thereby always reach their switching points between the two regimes at the same time. If, alternatively, the typical situation is that there are many firms facing both regimes and the share of firms facing each regime varies over time, the dynamic process is better described using continuous weights corresponding to the proportion of the firms in the aggregate that are under each regime in the different periods. In both cases, we estimate long-run solutions including relative prices, foreign demand and capital stock. The estimated price elasticities are quite similar in both models and consistent with what is reasonable to expect for relatively homogenous manufactures like metals.

## 2. Economic framework

We assume that a producer always faces two potential constraints on its exports, foreign demand and output capacity. As long as existing capacity allows for increased production, exports are determined by demand, given the price. When overall capacity utilisation rises, a switching point is reached where output capacity takes over as the export constraint. When operating at an aggregated level, this implies that all the producers reach their switching point at the same time. This is discussed more thoroughly later in this section. First, we take a closer look at a graphical illustration of the switching-mechanism.

**Figure 2.1.**




In figure (2.1), X* illustrates the switching point of a producer. When the output equals X*, the producer can not increase supply without increasing the output capacity. At this point, the marginal cost curve (MC) is vertical. When the output is below X*, as for instance $X_1$, the price is set by the producer and the output is determined by demand. Quantity and price are determined by the point where the marginal income curve ($MI_1$), derived from the demand curve, and the marginal cost curve intersect. If the demand curve (D) shifts to the left or to the right, exports will decrease or increase respectively. When the marginal income curve ($MI_2$) crosses the marginal cost curve in the vertical section, changes in demand no longer affect the quantity exported. Then, shifts in demand only affect the price, and exports are constrained by supply.

The producer can increase its supply by increasing output capacity. That will make the marginal cost curve, and thus X*, shift to the right. In economic literature, real capital is traditionally treated as fixed in the short run. In capital intensive industries, as the metal industry, extensive investments are required to increase output capacity. Such investments are expensive and take time to implement, and will therefore only take place when a positive shift in demand is expected to be permanent. Other factors that can come in the way of new investments are for instance a shortage of qualified workers or insufficient availability of inexpensive electrical power. When output capacity is fully utilised and the producer considers a shift in the demand curve to be temporary (and therefore chooses not to increase output capacity), or faces other conditions that keeps him from expanding capacity, the export constraint is on the supply side.

Based on this framework, we formulate an econometric model consisting of two separate parts, including demand variables in one part and supply variables in the other. We estimate two versions of this model. In the first version, the supply model is employed when output capacity is fully utilised, and the demand model is operating when spare capacity is available. Discrete weights, based on capacity utilisation-data, determine which model is operating in each period. Applying this model to aggregate data, i.e. data for total exports, we implicitly assume that all producers are under the same regime at all times and that they will jump instantaneously from one regime to the other in response to small changes in capacity utilisation. A more realistic motivation for this model is that if the same cycles apply to most of - or all - the firms in an aggregate, and if the firms' switching-points fall within a narrow range, we may get an acceptable approximation of the dynamic process by using discrete weights.

The alternative approach is to formulate the transition from one regime to the other as a gradual process. Instead of treating exports as determined either by demand or supply, it may in some cases be more realistic to assume that the aggregate consists of firms under both regimes at all times. The capacity utilisation-data is now used to construct continuous weights for the supply and demand variables. As the capacity utilisation increases, the supply model offers a relatively better description of exports and is given a larger weight. When the capacity utilisation decreases, the demand model becomes more important and thus is given a larger weight. Batchelor (1977) suggests an approach of this kind.

The export model with weights can be expressed as

$$(2.1)\quad a_t = \alpha_1 + D_t[\alpha_2 mii_t + \alpha_3 (pa - pmet)_t] + S_t \alpha_4 k_t + u_t ,$$

where lower case letters indicate logs. The elasticities of export (A) with respect to foreign demand (MII), relative prices (PA/PMET) and stock of capital (K) at time t are given by $\alpha_2$ and $\alpha_3$ multiplied by D, and by $\alpha_4$ multiplied by S, respectively. The signs of the coefficients are hypothesised as follows: $\alpha_2 > 0$, $\alpha_3 < 0$ and $\alpha_4 > 0$, meaning that exports should increase with foreign demand and stock of capital, and decrease with the Norwegian export price (PA) relative to the world market price (PMET). The equation is extended with an error term $u_t$. D and S are given two alternative interpretations, as discrete and continuous weights. In section 4.1, an export model with discrete weights is estimated. Then, D = 1 and S = 0 when exports are determined by demand, and D = 0 and S = 1 when exports are determined by supply. In section 4.2, D and S are defined as continuous weights. The discrete and continuous weights are based on an index for capacity utilisation in the Norwegian metal industry.

As it stands, Equation (2.1) may be interpreted as a static equilibrium, describing Norwegian exports of metals in the long run. Hence, it will serve as the foundation for the empirical analysis in section 4. Equation (2.1) may not hold in the short run, as it is likely that exports adjust gradually to changes in the explanatory variables. The empirical representation will therefore be specified dynamically.

If the market share of Norwegian exporters of metals is constant at constant relative prices, and there are no constraints on supply, the elasticity of exports with respect to demand equals one ($EL_{MII}A = 1$). This means that a one per cent increase in demand for metals in the world market leads to a one per cent increase in Norwegian exports of metals. Armington (1969) imposes this restriction. Market shares may be altered by factors such as marketing and relative changes in quality between competing products however. Such factors are difficult to observe, and are not included explicitly as explanatory variables in the econometric models presented here. We do not presuppose constant market shares at constant relative prises in the present paper, but the restriction is tested[1].

The export equation is restricted to be homogenous in prices, both in the short and in the long run. While price homogeneity in the long run is included because it is a theory restriction, price homogeneity in the short run is included to reduce the number of coefficients to be estimated. Rejection of price homogeneity in the short run may for instance be caused by asymmetric information or by contracts delaying the response to changes in prices. Rejection of long-run price homogeneity would indicate rejection of the long-run relationship, but could also be caused by measurement error. The price homogeneity restrictions are tested in section 4.

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[1] Goldstein and Khan (1985) conclude that income elasticities for industrial countries fall in the range of 1 to 2, and they find a higher elasticity for manufactures than for non-manufactures.

# 3. Data

The empirical analysis is conducted using quarterly data that spans the period of 1971(1)-1997(4). The sample period used for estimation is 1972(1)-1997(4) to allow for lags and transformations. Unless otherwise noted, data is taken from the Quarterly National Accounts, Published by Statistics Norway. See appendix 1 for detailed data definitions and sources. Figure 3 gives a description of the data.

**Figure 3.1. Historical development**[2]

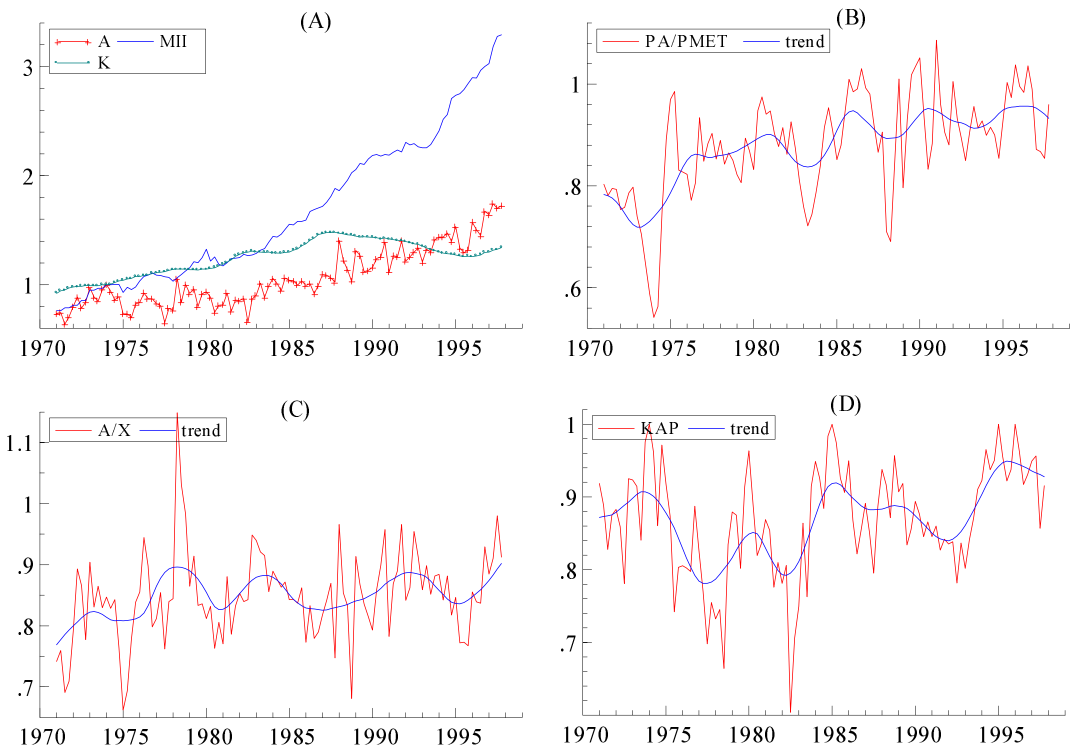


Figure 3(A) shows that export (A) is increasing at a slower rate than foreign demand (MII). This implies that the market elasticity is less than one, or that there are other factors pulling exports downwards. Among such factors, one could find for instance limited output capacity or an increase in relative export prices (PA/PMET). Figure 3(B) shows an increasing trend in PA/PMET lasting through the whole sample period. The export quantity variable, A, is derived by deflating a value series for Norwegian exports of metals by the export price index (PA). MII is an indicator for the demand for metals in the Norwegian export markets, and is calculated at Statistics Norway. PA is a price index for

[2] In figure 3(A), 3(B) and 3(D), all variables are normalised to one in 1974:1.

Norwegian exports of metals. As an indicator for the international price on metals, PMET, we use the International Monetary Fund's (IMF) price index for metals, measured in Norwegian kroner.

From the early 1980s, exports were increasing at a higher rate than the capital stock (K). Figure 3(C) shows that the relationship between export and output (X) is relatively stable throughout the eighties and the nineties, but increasing during the seventies. Therefore, we conclude that the increase in export relative to the capital stock is not due to an increase in the export share, i.e. the share of output being exported. The export-share averages 87 per cent in the sample.

An index for capacity utilisation in the Norwegian metal industry (KAP) is used to construct weights. Figure 3(D) shows the development of KAP over the sample period and indicates an increasing trend starting in the early 1980s. KAP is calculated by the Modified Wharton-Method (Cappelen and von der Fehr, 1986), based on data on output and capital stock for the Norwegian metal industry.

If the export volume and the export price are determined at the same time, use of PA as an explanatory variable may cause simultaneity bias in the coefficient estimates. Furthermore, when a quantity variable is derived by deflating an error free value series by a price variable subject to random errors, errors in the quantity variable are negatively correlated with the price variable. This is a well-known result, and Kemp (1962) has shown that the estimated export price elasticity will be biased towards -1. Hence, empirically estimated price elasticities that are larger, in absolute value, than -1 are too small, and elasticities that fall between 0 and -1 are too large in absolute value.

# 4. Empirical Results

We estimate two export equations, both based on equation (2.1). The specification of equation (2.1) implies that the full effect of changes in the exogenous variables on exports is achieved in the same quarter as the changes take place. Asymmetric information and adjustment costs prevent economic agents from instantaneous adjustment to new information. Econometric export equations should therefore be specified dynamically, and equation (2.1) is to be interpreted as a long-term relationship. As the lag-structure is not known, we start out by estimating a general dynamic version of equation (2.1), represented by the equilibrium-correction model in equation (4.1)[3].

[3] The empirical analysis was carried out using PcGive Version 9.2 [cf. Hendry and Doornik (1996)].

$$
\begin{aligned}
(4.1)\qquad \Delta a_t &= \alpha + \sum_{i=1}^{5} \beta_{1t-i} \Delta a_{t-i} + \lambda_1 a_{t-1} \\
&+ \sum_{i=0}^{5} D_{t-i} \beta_{2t-i} \Delta mii_{t-i} + \sum_{i=0}^{5} D_{t-i} \beta_{3t-i} \Delta (pa - pmet)_{t-i} + D_{t-1} [\lambda_2 mii_{t-1} + \lambda_3 (pa - pmet)_{t-1}] \\
&+ \sum_{i=0}^{5} S_{t-i} \beta_{4t-i} \Delta k_{t-i} + S_{t-1} \lambda_4 k_{t-1} \\
&+ \sum_{q=1}^{3} DUM_q + u_t
\end{aligned}
$$

Lower case letters indicate logs. $D_{t-i}$ and $S_{t-i}$ are weights for the demand and supply variables respectively. The weights are defined as discrete and continuous in the analyses in section (4.1) and (4.2) respectively. Δ is the difference operator, $DUM_q$, q = 1,2,3, are seasonal dummy variables and $u_t$ is the error term assumed to be white noise.

Price homogeneity is implemented in the general model. Short-run price homogeneity is tested by including $D_t \Delta pa_t$ in addition to the variable $D_t \Delta(pa\text{-}pmet)_t$ [$=D_t(pa\text{-}pmet)_t$ - $D_t(pa\text{-}pmet)_{t-1}$ = $D_t \Delta pa_t$ - $D_t \Delta pmet_t$]. Price homogeneity in the long run is tested by including $D_{t-1}pa_{t-1}$ in addition to the variable $D_{t-1}(pa\text{-}pmet)_{t-1}$. The t-values for the included variables are used as test statistics. These test statistics are presented together with the other statistic tests following equation (4.2) and equation (4.4). If $\Delta pa_t$ or $pa_{t-1}$ is significant, price homogeneity in the short and long run respectively is rejected.

Our modelling strategy is the general to specific approach [cf. Davidson et al. (1978)]. First, insignificant parameters are set equal to zero. Then, variables with coefficients with signs not consistent with theory are rejected. The model is also made more general based on various reduced variants to test the robustness of earlier restrictions. Restrictions between parameters are implemented if it contributes to simplify the model without reducing the fit (measured by the equation standard error). We also emphasise that the final estimated equation should pass standard tests for autocorrelation, heteroscedasticity and normality.

## 4.1. An export relation with discrete weights

First, we focus on a model with discrete weights, assuming that this specification is acceptably close to the underlying or "true" export relationship. $D_{t-i}$ and $S_{t-i}$ are defined as discrete weights for period t-i, i = 0,1,.,5. D = 1 and S = 0 when exports are determined by demand and D = 0 and S = 1 when exports are determined by supply. A 95 per cent switching point was chosen as a starting point for the analysis: The supply model is used when capacity utilisation is 95 per cent or more. When capital utilisation is less than 95 per cent, the demand model is employed. A sensitivity analysis of the choice of switching point is carried out later in this section. The general to specific approach leads to the equilibrium-correction model in equation (4.2). The standard error is given in the parentheses under each estimate. DUMiq are dummy variables for outliers in quarter q, year i.

$$\begin{aligned}
\Delta a_t &= \underset{(0.11)}{0.35} - \underset{(0.07)}{0.30}\Delta a_{t-1} - \underset{(0.11)}{0.99} D_t \Delta pa_t + \underset{(0.10)}{0.33} D_t\, \Delta pmet_t + \underset{(0.27)}{1.04} D_t \Delta mii_t \\
(4.2) \qquad & - \underset{(0.08)}{0.33} D_{t-1}(pa - pmet)_{t-1} - \underset{(0.03)}{0.09}(a - Dmii - 0.5Sk)_{t-1} \\
& + \underset{(0.01)}{0.03} DUM_1 + \underset{(0.01)}{0.08} DUM_3 + \underset{(0.06)}{0.27} DUM782 - \underset{(0.06)}{0.19} DUM884 + \underset{(0.06)}{0.17} DUM961
\end{aligned}$$

Method: OLS 1972:1 - 1997:4 $R^2 = 0.75$ F(11,92) = 25.0 [0.0000] $\sigma = 0.06$ DW = 2.26
RSS = 0.33 for 12 variables and 104 observations

| Residual misspecification tests: | | | |
|---|---|---|---|
| AR 1- 5 | F( 5, 87) = | 1.83 [0.11] | |
| ARCH 4 | F( 4, 84) = | 0.10 [0.98] | |
| Normality | $\chi^2(2)$ = | 0.87 [0.65] | |
| $X_i^2$ | F(18, 73) = | 0.57 [0.91] | |
| $X_i \cdot X_j$ | F(45, 46) = | 0.73 [0.85] | |
| RESET | F( 1, 91) = | 0.004 [0.95] | |

Tests of restrictions:

| | |
|---|---|
| $D\Delta pa_t$: | t-value =-4.42 |
| $Dpa_{t-1}$: | t-value = 0.57 |
| $Dmii_{t-1}$: | t-value = 0.47 |
| $Sk_{t-1}$: | t-value = 0.55 |

$R^2$ is the squared multiple correlation coefficient, F is an F-test of whether $R^2$ equals zero, $\sigma$ is the residual standard error, DW is the Durbin-Watson statistic and RSS is the residual sum of squares. In addition we report a testing sequence on the residuals for a range of null hypotheses of interest, including no autocorrelation (AR 1-5), no heteroscedasticity (ARCH, $X_i^2$ and $X_i \cdot X_j$), normality of the distribution of the residuals, and no functional form misspecification[4]. None of the tests of the residuals are significant.

Parameter constancy is studied by recursive estimation. The plots A-I in Appendix 2 are of the coefficient estimates at each point in the sample together with their 95 per cent confidence intervals. For most of the data set, the estimates are significantly different from zero, and are fairly constant over the sample. Plot J shows the residuals from the recursive estimation with their 95 per cent confidence intervals. The confidence bands are relatively constant, and none of the residuals are outside the bands. Residuals outside the bands would indicate outliers. Parameter constancy is not rejected by the break point Chow test [see Chow (1960)] in the final plot.

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[4] See Hendry and Doornik (1996) p.237-241 for a description of these tests.

The static long-run equilibrium for equation (4.2) is

$$(4.3)\quad a = 3.80 + D{\cdot}mii - 3.68{\cdot}D{\cdot}(pa - pmet) + 0.5{\cdot}S{\cdot}k$$

The coefficients are interpreted as partial long-run elasticities of exports with respect to foreign demand (MII), relative prices (PA/PMET) and capital stock (K). The estimated long-run price elasticity is considerably higher than what have been found in previous studies of Norwegian export of metals[5]. It is now estimated to be -3.68 in periods when the export price is set by the supply side and exports are given by demand. This estimate is more in line with our prior expectation for relatively homogenous manufactures like metals. When the export constraint is on the supply side, the price elasticity equals zero. Given the 95 per cent switching point, the demand variables MII, PA and PMET explain exports in about 83 percent of the periods in the sample. In the remaining periods, exports are determined by K.

Price homogeneity is rejected in the short run, but not in the long run [see the results of the tests of the restrictions below equation (4.2)]. Agreements or contracts, delaying response in the market to changes in prices, may explain that price homogeneity is a valid restriction in the long run only. Another explanation could be that it takes some time before changes in relative prices are identified because of asymmetrical information.

Because of improved performance in statistical terms, we apply restrictions on the long-run effects of foreign demand (MII) and capital stock (K) on exports. Only restrictions supported by the data are included. The market elasticity is set to one ($EL_{MII}A = 1$). The elasticity of exports (A) with respect to capital stock (K) is set to 0.5. These restrictions are tested, and not rejected, using the same method as in the case of price homogeneity, by including $Dmii_{t-1}$ and $Sk_{t-1}$ respectively in addition to $(a\text{- }Dmii - 0.5Sk)_{t-1}$ [see equation (4.2)]. These restrictions are also consistent with the graphic analysis in section 3.

As pointed out in section 3, the estimated export price elasticity will be biased towards -1 when the quantity variable is derived by deflating an error free value series by a price variable subject to random errors. In equation (4.2), the short-run export price elasticity is estimated to be -0.99 and is significant

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[5] In both Lindquist (1993) and Naug (1995a), the estimated price elasticities are below one in absolute value.

with a t-value of -8.82. If this is a result of erroneously measuring changes in the export price, the residual standard error and thus the variable standard errors are underestimated. This would lead to overvalued t-values. Lindquist (1993) estimates export volume relations for the Norwegian metal industry using instrumental variables for the export price. She concludes that the use of instrumental variables has very little influence on the results.

A 95 per cent switching point was chosen as a starting point for this analysis. The switching point is decided upon at an a priori basis. Therefore, the robustness of the results to the choice of switching point deserves a closer investigation. A sensitivity analysis (see table 4.1) shows that the model is relatively robust to small changes in the switching point. When the switching point is chosen between 93 and 97 per cent, the coefficient estimates are stable, none of the variable restrictions are rejected and the equation standard errors stay at an acceptable level. With a switching point at 93 per cent and below, however, the hypothesis that the residuals have a normal distribution is rejected at a one per cent level, and all the long-run variables in the model are rejected. The share of the sample periods where the supply model applies range from 28 percent when a 93 per cent switching point is used, to 8 per cent using a 98 per cent switching point, getting disturbingly low at the higher range. The conclusion is that the results are quite robust when the choice of switching point falls within the rather narrow range of 94 to 97 per cent.

**Table 4.1. Sensitivity Analysis of the Switching Point[a]**

| Switching point | $\hat{\alpha}_0$ | $\hat{\alpha}_1$ [b] | $\hat{\alpha}_2$ [c] | $\hat{\alpha}_3$ [d] | Number (share) of periods supply model applies | Rejected variables[e] | Equation standard error | Diagnostic tests[f] |
|---|---|---|---|---|---|---|---|---|
| 98 | 3.75 | 1 | 4.05 | 0.5(R) | 8 (7.69%) | | σ = 0.057 | |
| 97 | 3.81 | 1 | 3.55 | 0.5 | 12 (11.54%) | | σ =0.059 | |
| 96 | 3.81 | 1 | 3.51 | 0.5 | 17 (16.35%) | | σ = 0.060 | |
| 95 | 3.80 | 1 | 3.68 | 0.5 | 19 (18.27%) | | σ = 0.060 | |
| 94 | 3.83 | 1 | 3.73 | 0.5 | 23 (22.12%) | | σ = 0.063 | Normality* |
| 93 | 3.79 | 1 | 4.24 | 0.5 | 29 (27.88%) | | σ = 0.067 | Normality** |
| 92 | 3.69 | 1(R) | 6.31(R) | 0.5(R) | 38 (36.54%) | All long-run variables | σ = 0.071 | Normality** |
| 90 | 3.61 | 1(R) | 9.70(R) | 0.5(R) | 46 (44.23%) | All long-run variables | σ = 0.073 | Normality** |

[a] The estimated long-run model (with a 95 per cent switching point): $a = \hat{\alpha}_0 + \hat{\alpha}_1 Dmmi + \hat{\alpha}_2 D(pa - pmet) + \hat{\alpha}_3 Sk$

[b] Restricted to one à priori. (R) means restriction is rejected.

[c] (R) means long-run price homogeneity is rejected.

[d] Equals 0.5 per restriction. (R) means restriction is rejected.

[e] Compared to the model with a 95 per cent switching point.

[f] The following null hypotheses are tested: No autocorrelation, no autoregressive conditional heteroscedasticity, normality of the distribution of the residuals, no heteroscedasticity and no functional form misspecification (see Hendry and Doornik (1996), p. 50 and 237-241). If a null hypothesis is rejected, we specify which and whether it is rejected at a five (*) or one (**) per cent level.

## 4.2. An export relation with continuous weights

The model with discrete weights depends on the non-trivial assumption that all producers reach their switching point at the same time. This implies that all producers are under the same regime at all times and, as a consequence, that they all reach their capacity limit at the same time. In this section, we propose an alternative approach by postulating that at any one time the firms experience a variety of capacity utilisation conditions, which are generally high at peaks of the business cycle and low in troughs. Instead of considering export as determined by either demand when price is set or by supply when capacity is fully utilised, both demand and supply variables are included in the model at the same time. The index for capacity utilisation in the Norwegian metal industry (KAP) is now used to construct continuous weights for the supply and demand variables. As capacity utilisation increases, the supply variables are assumed to offer a relatively better description of export, and the supply model is given a larger weight. As capacity utilisation decreases the demand model becomes more important and thus is given a larger weight [see Batchelor (1977) and Benedictow (1999)].

Before we estimate the model, we have to decide how to construct the continuous weights. The approach here is to scale the weights for the supply and demand variables, S and D respectively, in such a way that both equal one on average. The average value of the capacity utilisation index, KAP, is 88 per cent. S is calculated by multiplying the KAP-series by 1.14 (= 1/0.88). The weight for the demand model, D, is calculated as follows: D = 1 - (S - 1) = 2 - S. This makes the scaling of KAP neutral in the sense that it does not alter the initial relative importance of supply and demand variables on average. S (D) is larger (smaller) than one in periods where KAP takes on a value higher than its own average, and smaller (larger) than one when KAP is smaller than 0.88. S varies between 0.69 when capacity utilisation is at its lowest level and 1.14 at full capacity utilisation. D varies correspondingly between 1.31 and 0.86. The general to specific approach now leads us to the equilibrium-correction model in equation (4.4).

$$\begin{aligned}(4.4)\qquad \Delta a_t &= \underset{(0.03)}{0.08} - \underset{(0.06)}{0.32}\,\Delta a_{t-1} - \underset{(0.06)}{0.16}\,\Delta a_{t-3} + \underset{(0.25)}{1.16}\,D_t \Delta mii_t - \underset{(0.11)}{1.02}\,D_t \Delta pa_t + \underset{(0.09)}{0.30}\,D_t \Delta pmet_t \\ &\quad - \underset{(0.07)}{0.41}\,D_{t-1}(pa - pmet)_{t-1} - \underset{(0.03)}{0.14}(a - Dmii - \tfrac{1}{3}Sk)_{t-1} \\ &\quad + \underset{(0.01)}{0.04}\,DUM_1 - \underset{(0.01)}{0.07}\,DUM_3\end{aligned}$$

Method: OLS 1972:1 - 1997:4 $R^2 = 0.75$ $F(9,94) = 31.11$ [0.00] $\sigma = 0.06$ DW = 2.15
RSS = 0.33 for 10 variables and 104 observations

Residual misspecification tests:

| | | | |
|---|---|---|---|
| AR 1- 5 | F( 5, 89) | = | 0.91 [0.48] |
| ARCH 4 | F( 4, 86) | = | 0.68 [0.61] |
| Normality | $\chi^2(2)$ | = | 6.70 [0.0352] * |
| $X_i^2$ | F(17, 76) | = | 0.45 [0.97] |
| $X_i \cdot X_j$ | F(52, 41) | = | 0.47 [0.99] |
| RESET | F( 1, 93) | = | 0.02 [0.90] |

Tests of restrictions:

| | |
|---|---|
| $D\Delta pa_t$: | t-value = -5.45 |
| $Dpa_{t-1}$: | t-value = -0.73 |
| $Dmii_{t-1}$: | t-value = -0.65 |
| $Sk_{t-1}$: | t-value = 0.10 |

The model performs reasonably well confronted with the residual misspecification tests, although there is indication of non-normality. The normality test is significant at the 5 per cent level (*), but not at the 1 per cent level. The equation standard error is very similar to what we found in the model with discrete weights.

Again, recursive estimation is used to study parameter constancy. The parameter estimates are relatively constant over the sample and parameter constancy is not rejected by a break-point Chow test. The confidence bands for the residuals are fairly constant and there are no strong indications of outliers (see appendix 3).

The static long-run solution is given by

(4.5) $a = 0.56 + D{\cdot}mii - 3.02{\cdot}D{\cdot}(pa - pmet) + 0.33{\cdot}S{\cdot}k$

The estimated price elasticity is -3.02·D. This implies that the elasticity varies between different periods, depending on the level of capacity utilisation represented by D, but takes on an average value of -3.02. This is close to the estimated price elasticity in the model with discrete weights, which is -3.68 in periods where D equals one (in 83 per cent of the periods)[6]. Like in section 4.1, short-run price homogeneity is rejected, and the market elasticity is restricted to average unity in the long run. The elasticity of the capital stock is restricted to 0.33 (on average), compared to 0.5 in periods of full capacity utilisation when discrete weights are employed. These restrictions are supported by the data.

The short-run export price elasticity is estimated to be -1.02 and is significant with a t-value of -9.09. This could, as in the model with discrete weights, be a result of simultaneity between A and PA, and the t-values may for the same reason be overvalued.

## 5. Conclusions

The purpose of this analysis is to test two alternative models for export determination, both within a framework of differentiated products and limited output capacity. Demand and supply variables are weighted in accordance to their relative importance in each period. A measure of capacity utilisation is used to construct discrete and continuous weights. We estimate export equations including foreign demand, relative prices and capital stock. We are not able to discriminate between the model with discrete weights and the model with continuous weights on statistical ground. The estimated elasticities are very similar, as well as the residual standard errors, and both models perform well confronted with several tests for misspecification and parameter stability.

The estimated long-run price elasticities are considerably higher (in absolute value) than what have been found in previous studies of Norwegian exports of metals. Both Lindquist (1993) and Naug (1995a) find estimates well below 1 in absolute value. In the model with discrete weights, the price elasticity is now estimated to be -3.68 in periods where exports are determined by demand given the export price. When the export constraint is on the supply side, price changes do not affect exports. In

[6] Hence, the corresponding average estimated price elasticity in the model with discrete weights would be -3.05 = -3.68·0.83.

the model with continuous weights, the estimated elasticity averages -3.02. These estimates are more consistent with what is reasonable to expect for homogenous manufactures like metals.

In both models, price homogeneity is rejected in the short, but not in the long run. Agreements or contracts, delaying response in the market to changes in prices, may explain that price homogeneity is a valid restriction in the long-run only. An alternative explanation is that it takes time before changes in relative prices are identified by the "consumers" because of asymmetrical information. Because of improved performance in statistical terms, we apply restrictions on the long-run effects of foreign demand and capital stock on exports. The market elasticity is restricted to unity in both models, and the elasticity of the capital stock is restricted to 0.5 and 0.33 in the models with discrete and continuous weights respectively. These restrictions are tested and not rejected.

Recursive estimation is used to study parameter constancy. In both models, the parameter estimates are relatively constant over the sample, and parameter constancy is not rejected by a break-point Chow test. The confidence bands for the residuals are fairly constant and there are no strong indications of outliers. Both models performs well confronted with a testing sequence on the residuals for a range of null hypotheses of interest, including autocorrelation, normality of the distribution of the residuals, heteroscedasticity and functional form misspesification, although there is indication of non-normality in the model with continuous weights. In the model with discrete weights, the switching point is decided upon at an a priori basis. Therefore, the robustness of the results to the choice of switching point is subject to a closer investigation. A sensitivity analysis shows that the model is relatively robust to small changes in the switching point.

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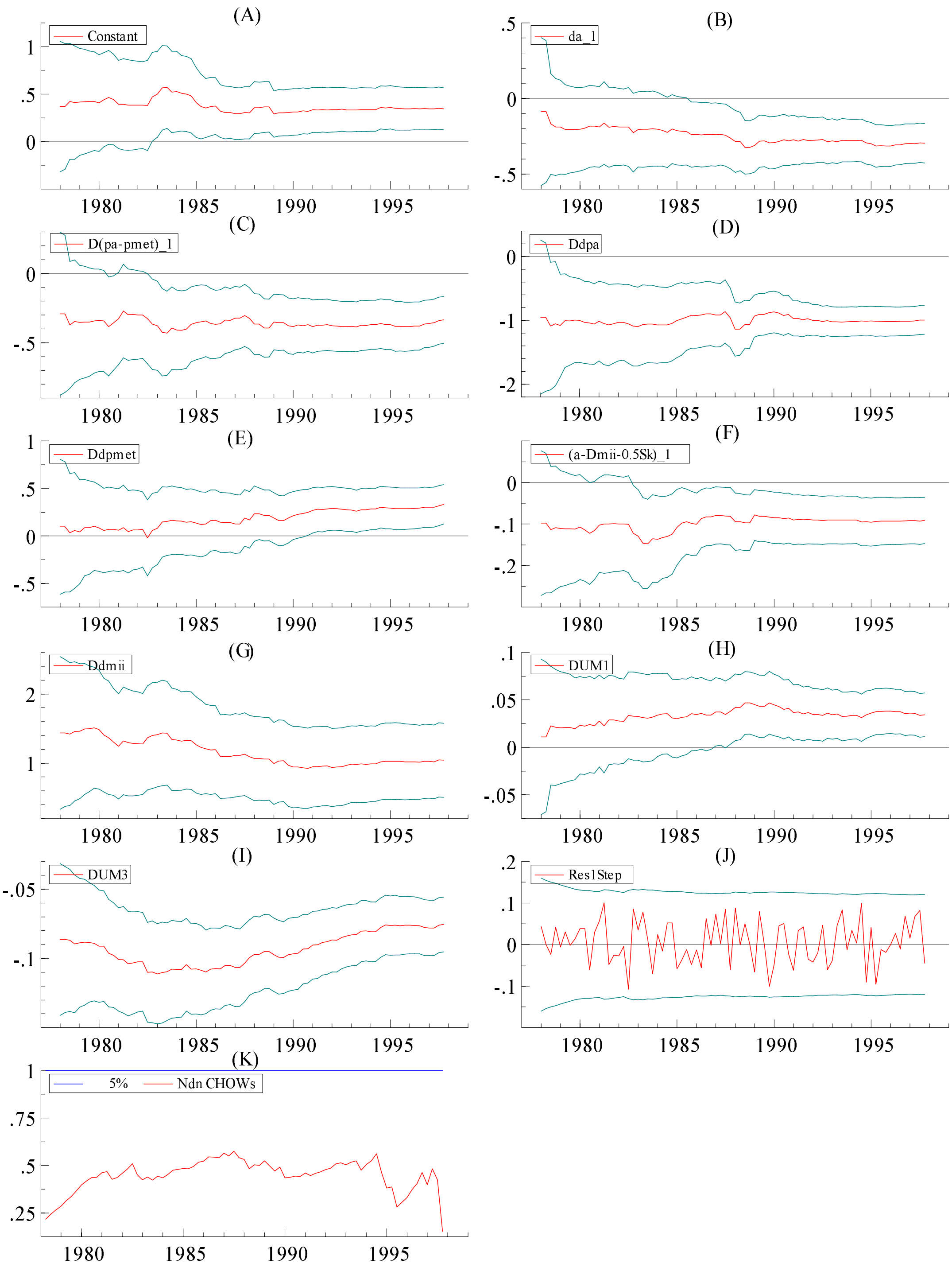
**Appendix 1**

**Data Definitions and sources**

| | |
|---|---|
| A | Norwegian exports of metals in constant prices. Source: Statistics Norway. |
| MII | Weighted volume index for total commodity exports demand faced by Norwegian manufacturing. Source: Statistics Norway. |
| PA | Price index for Norwegian exports of metals. Source: Statistics Norway. |
| PMET | Price index for metals (measured in Norwegian kroner). Source: The International Monetary Fund (IMF). |
| K | The stock of capital in the Norwegian metal industry. Source: Statistics Norway. |
| KAP | Index for capital utilisation in the Norwegian metal industry. Source: Statistics Norway. |
| DUMiq | Dummy variable used to account for an outlier in the equation. Equals 1 in year i, quarter q, 0 otherwise. |
| $DUM_q$ | Seasonal dummy variable for quarter q, equals 1 in quarter q, 0 otherwise, i = 1,2,3. |

# Appendix 2

**Equation (4.2): Recursive least squares graphical constancy statistics. d = Δ.**



**Appendix 3**

**Equation (4.4): Recursive least squares constancy statistics. d = Δ.**