**Discussion Papers No. 243, January 1999**
**Statistics Norway, Research Department**

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# Modelling the Demand for Imports and Domestic Output

**Abstract:**
The paper models domestic output over imports in Norway's expenditure on manufactures. Using Johansen's (1988, 1991) method, we obtain a cointegrating vector between the output-imports ratio, relative prices and a proxy for international specialisation. This vector enters a conditional equilibrium correction model of the output-imports ratio; a model which also includes short-run influences of relative prices and a negative coefficient for domestic capacity utilisation. The utilisation coefficient aside, we do not find significant activity effects on the output-imports ratio. Lastly, the model passes several tests of the Lucas critique.

**Keywords:** Import share, international specialisation, cointegration, Johansen procedure, equilibrium correction model, parameter constancy, Lucas critique.

**JEL classification:** C22, C32, F12, F14.

**Acknowledgement:** Comments from Å. Cappelen, K. Moum, R. Nymoen and F. Steen are gratefully acknowledged. The results were obtained using PcGive Professional Versions 9.0 and 9.1; cf. Hendry and Doornik (1996) and Doornik and Hendry (1996, 1997).

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# 1. Introduction

The allocation of domestic demand between imports and home goods has important implications for policy. Perhaps most important, the effectiveness of demand policies depends critically on import shares and their relation with domestic demand. The price sensitivity of imported and domestic goods, likewise, may be vital for the outcome of exchange rate changes. It is unsurprising therefore that many studies have estimated equations for imports or import shares (see e.g. Goldstein and Khan's (1985) survey). There is a case for more research in this field, however, as previous import studies are subject to several limitations.

First, most of them report long-run activity elasticities which are prone to omitted-variables bias. Digressing briefly, the standard approach is to estimate conditional import models with domestic activity and relative prices as regressors. Such models typically yield long-run activity elasticities considerably above one, implying that import shares increase with domestic demand. Import shares in OECD countries have indeed shown an upward trend since 1950; but this trend may be due to increasing trade specialisation, driven by *supply* factors (cf. e.g. Krugman (1995)). If so, and if the specialisation effects are not picked up by measured relative prices, conventional import equations overstate the response of imports to demand changes. Moreover, a number of import functions in the literature include domestic capacity utilisation or related variables. As shown by Barker (1979), long-run activity effects are likely to be overvalued in import equations which (a) wrongly omit capacity utilisation, but (b) correctly include activity and a trending proxy for supply influences. Yet, to our knowledge, only a study on Austrian data and four UK studies have included effects of activity, specialisation and capacity utilisation in import (share) models.[1] Except for Cuthbertson (1985), they obtain unit long-run elasticities of imports with respect to activity.

Second, the early contributions and many recent studies disregard non-stationarity in the data, use static or restrictive dynamic specifications, and fail to test properly for residual misspecification, weak exogeneity and parameter stability. It cannot be denied that econometric practice has improved greatly since the advent of cointegration theory and the LSE methodology. However, import papers incorporating these advances often use Engle and Granger's (1987) static-regression procedure and/or estimate conditional models without testing for weak exogeneity. The former approach can result in severely biased long-run estimates and cointegration tests with low power (cf. Banerjee *et al.* (1986)

and Kremers *et al.* (1992)); and conditional analysis is valid only if the conditioning variables are weakly exogenous for all parameters of interest,[2] an assumption which is far from obvious.[3] The Johansen (1988, 1991) method goes a long way towards resolving these problems, but Johansen's methodology has not been widely used in import studies. (See, however, Sedgley and Smith (1994), Urbain (1995, 1996), Menon (1995), Carone (1996) and Abbot and Seddighi (1996)).

Third, and relatedly, import modellers have paid insufficient attention to the Lucas (1976) critique.[4] Specifically, most studies simply assume that (i) the import demand parameters are invariant to policy interventions; and (ii) purchasers do not act on model-based expectations of prices and activity. Both assumptions are questionable in a world of imperfect information and adjustment costs. Hence some papers (such as Husted and Kollintzas (1987), Gagnon (1989), Clarida (1994, 1996) and de la Croix and Urbain (1996)) have modelled imports within a rational expectations framework. These papers do not check whether competing models suffer from Lucas's critique, however.[5] In our view, that hypothesis should be tested rather than postulated.

Using Norwegian data over 1967(1)–1994(4), we conduct an import share study that seems to overcome these limitations. The ratio of domestic manufacturers' home sales to competing imports – denoted "the output-imports ratio" – is modelled by a specification which allows for relative prices, domestic activity, capacity utilisation and a proxy for trade specialisation. Applying Johansen's method to the data, we obtain a cointegrating vector for the output-imports ratio with effects of specialisation and relative prices. Next, taking the estimated cointegration parameters as known, we develop a conditional equilibrium correction model of the output-imports ratio. This model passes a battery of tests for residual misspecification, weak exogeneity, parameter stability and the Lucas critique. In addition to disequilibrium correction effects, it contains short-run influences of relative prices and a negative coefficient for capacity utilisation. As could be expected from our arguments above, we do not find significant effects of the activity variable.

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[1] See the equation for Austrian imports in Neudorfer *et al.* (1990, Table 3.7) and the UK studies by Cuthbertson (1985), Anderton *et al.* (1992), Sedgley and Smith (1994) and Temple and Urga (1997).

[2] See Engle *et al.* (1983), Hendry and Neale (1988), Johansen (1992) and Hendry (1995a).

[3] Conventional theory notwithstanding, import prices in small countries may well respond to domestic demand changes (see Naug and Nymoen (1996)); activity variables could be endogenous by including imports. Finally, weak exogeneity of prices and activity is precluded if, as discussed below, agents act on expectations models of those variables; cf. Hendry and Neale (1988).

[4] Exceptions are Urbain (1995) (using Belgian data) and Carone (1996) (on US data). Also, inspired by the classic Orcutt (1950) paper, Liu (1954), Goldstein and Khan (1976) and Wilson and Takacs (1979) test claims that parallel the Lucas critique. The general lack of testing is documented by Ericsson and Irons (1996).

[5] An expectations model alone, however congruent it may appear, is not compelling evidence of the Lucas critique; see e.g. Hendry (1988, pp. 135-36) and Ericsson and Irons (1995, p. 267).

## 2. Economic Background

Following standard practice in import studies, home-produced goods and competing imports are assumed to be imperfect substitutes. We further assume that domestic and imported manufactures form a separable branch of agents' utility/production functions; that this branch can be represented by a non-homothetic CES function (cf. Sato (1975, 1977)); and that agents maximise utility/profits. In the long run, their expenditure on manufactures (*Z*) is allocated between imports and home goods according to the following demand equation:

$$(1) \quad hi = \mu_0 + \mu_1 rp + \mu_2 z + \mu_3 s + \mu_4 CU.$$

Lower case variables are in logs; $HI = H/I$, where *H* is real domestic output sold in Norway and *I* is real imports including tariffs; $RP = PI/PH$, where *PI* is the tariff-adjusted import price in domestic currency and *PH* is the price of *H*; $Z = H+I$; *S* is a proxy for international specialisation; and *CU* is the rate of capacity utilisation in domestic manufacturing. The latter variable enters (1) linearly, since a log specification would imply (implausibly) that rises in *CU* decrease *hi* more at low levels of utilisation than at high levels (see the discussion below). Thus $\mu_4$ is a semi-elasticity, while $\mu_1$, $\mu_2$ and $\mu_3$ are the partial elasticities of *HI* with respect to *RP*, *Z* and *S*. Especially, $\mu_1$ is the long-run elasticity of substitution between home goods and competing imports.[6] The data series are described in Section 3, and Sections 4 and 5 obtain estimates of the coefficients.

The output-imports ratio *HI* and the import share *IMP* are related through the definition $IMP \equiv I/(I+H) \equiv 1/(1+HI)$. As an alternative to (1), we could have used Deaton and Muellbauer's (1980) AIDS specification to model *IMP* directly. AIDS models may predict $IMP > 1$, however, whereas equation (1) is data admissible in this respect (cf. Wallis (1987)). The admissibility issue aside, it proved difficult to discriminate empirically between (1) and the AIDS specification. Both functional forms have been employed in the imports literature; see, among others, Gregory (1971), Barker (1977, 1979), Lächler (1985), Anderton *et al.* (1992) and Sedgley and Smith (1994).

Turning to the role of the right-hand-side variables in (1), *z*, *rp* and *s* are likely to have operated in such a way as to decrease *hi* historically. In line with the high activity elasticities of many import models, Barker's (1977) variety hypothesis suggests that $\mu_2 < 0$. The idea is that consumers demand

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[6] Although theoretically plausible, long-run price homogeneity is not obvious empirically; see Murray and Ginmann (1976) and Urbain (1992a, 1995). However, as *pi* and *ph* might be I(2), the Johansen analysis does not allow for price inhomogeneity. Assuming *pi* and *ph* to be I(1), long-run homogeneity is not rejected for the model in Section 5.

more product varieties as their income increases; and, moreover, that economies of scale lead domestic firms to produce few varieties. Then, if the domestic varieties are bought by many domestic purchasers, as is often the case, one would expect *hi* to decrease with *z*. Next, liberalisations of trade flows and technological advances in transport have lowered real trade costs over time; a process which (*ceteris paribus*) has decreased *rp*.[7] Its effect on *hi* should also be captured by *s*, however, since trade costs make some foreign goods prohibitively expensive: lower trade costs for those goods may well trigger imports without changing *rp*. The *s* term should further pick up trade liberalisations unrelated to trade costs (although, strictly speaking, this precludes (1) from being a pure demand function). Lastly, it ought to reflect recent innovations in communication technology, the argument being that they have spurred trade by lowering natural trade barriers. In particular, it appears that the development of telecoms and computer systems – combined with the other supply influences – has driven a trend towards locating different parts of production processes in different countries; see Krugman (1995) and the survey in *The Economist*, September 28th 1996.

Capacity utilisation *CU* proxies for the extent of capacity constraints in domestic manufacturing. According to neoclassical theory, such constraints influence demand through prices only ($\mu_4 = 0$). In practice, however, firms have imperfect information about the state of goods markets and the effects of price changes. This might in turn cause price inertia,[8] inducing markets to clear by movements in non-price factors as well as by price changes (cf. Carlton (1989)). Specifically, increased presence of bottlenecks may yield less advertising, longer delivery times, less active pursuit of new orders and reduced ability to meet complex demands – each of which should lower the demand for domestic goods relative to competing imports.[9] The output-imports ratio is assumed to be independent of foreign capacity constraints, but the country-composition of imports need not be.

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[7] Both reductions in tariff rates and non-tariff barriers have lowered real trade costs. So have advances in cargo-handling, the evolution of air freight and the increased size of merchant ships; see van Bergeijk and Mensink (1997, p. 162), *The Economist*, November 15th 1997, pp. 89-90 and the comments by Cooper and Srinivasan in Krugman (1995).

[8] See e.g. Greenwald and Stiglitz (1989) and Andersen (1994, Ch. 6). Price inertia may also arise from costs of changing prices.

[9] The interpretation draws on Gregory (1971) and the UK studies cited in footnote 1. Alternatively, one may interpret $\mu_4 < 0$ within a disequilibrium framework; see the papers in Drèze and Bean (1990).

Owing to habits, contracts, delivery lags, adjustment costs and incomplete information, agents are expected to adjust *gradually* to changes in *rp* and *CU*. As the actual adjustment paths are unknown a-priori, the lag structure of the below model is determined from the data. The single equation analysis also allows for separate short-run coefficients of *pi* and *ph*: purchasers often have better information on some prices than on others, and so they may respond asymmetrically to changes in *pi* and *ph*. In addition, they may act upon data-based (as opposed to *model*-based) expectations, and data-based predictors of *pi* and *ph* need not have the same form. See Hendry and Ericsson (1991, Section 4.3) on data-based predictors and Urbain (1992a, 1996) on short-run inhomogeneity in import models. Section 6 addresses the issue of expectations formation.

# 3. The Data

In the empirical analysis, we use quarterly seasonally unadjusted data for the period 1967(1)-1994(4).[10] Allowing for five lags, the full sample estimations are over 1968(2)–1994(4). The Appendix gives details of the data series and their sources.

The output-imports ratio *HI* is defined as domestic manufacturers' home sales *H* over manufacturing imports (including tariffs) *I*. Both *H* and *I* exclude ships, oil platforms and the bulk of industrial commodities, however: sales of ships and platforms are erratic (and hence difficult to model), and the assumption of imperfect substitution is doubtful for many commodities. The *I* variable also omits non-competing imports (e.g. cars and planes). Import prices and domestic prices are measured by the implicit deflators of *I* and *H*.

The *CU* series is a weighted average of three disaggregated proxies for capacity utilisation. Each of them is estimated by a variant of the Klein and Summers (1966) "trend-through-peaks" (Wharton) method, with peaks in the output-capital ratio defining full utilisation (= 1). It is recognised that Wharton-type indicators are subject to deficiencies (cf. e.g. Perry (1973, pp. 708-09)), but our proxy appears to capture salient features of Norwegian business cycles (see the discussion of Figure 1 below).[11] Finally, the specialisation variable *S* is an eight-quarter moving average of manufacturing exports to manufacturing production in OECD countries (excluding Norway and recent OECD

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[10] Because of a change of calculation procedures in the national accounts, we were unable to use long and consistent series extending beyond 1994(4).

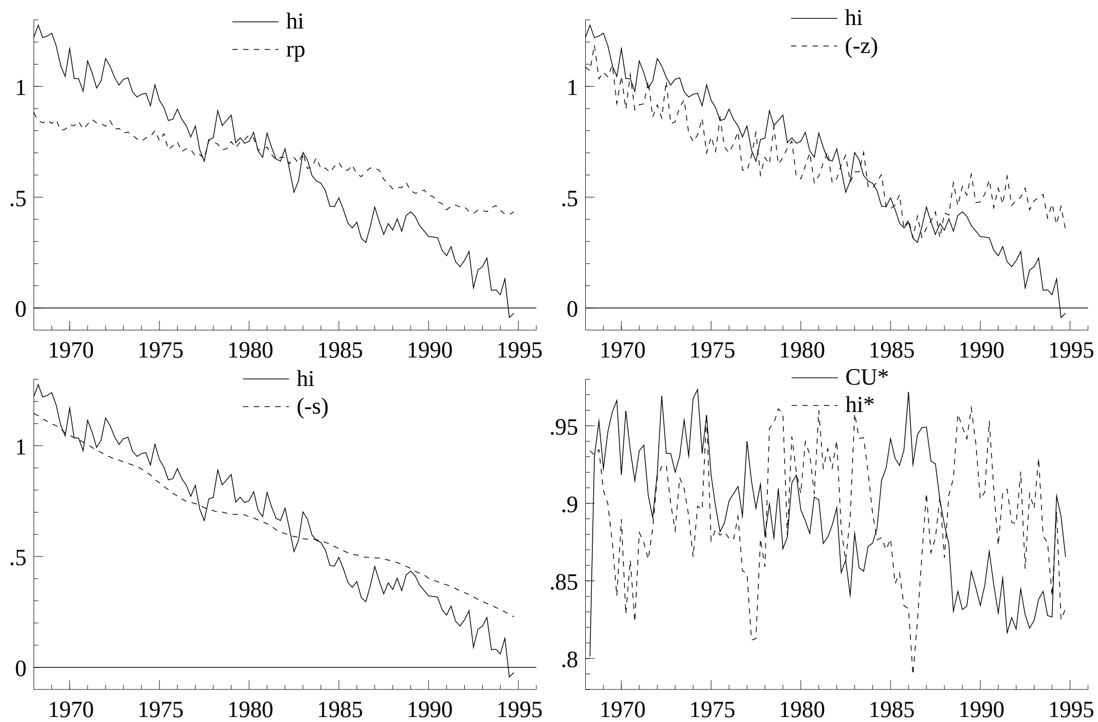
[11] Many import studies have used utilisation indicators based on business survey data. We did not explore this possibility, as survey data on utilisation are unavailable prior to 1973(4).

entrants). Similar specialisation measures have been employed by Beenstock and Warburton (1982), van Bergeijk and Mensink (1997) and the papers cited in footnote 1.

Except for *S*, these variables are constructed using data from the quarterly national accounts, QNA. Due to a change of calculation routines, most QNA series obtain a permanent shift in seasonality after 1977(4). To account for seasonality, therefore, we use three centred seasonal dummies ($Q_{1t}$, $Q_{2t}$ and $Q_{3t}$) and a step dummy $D_t$, equalling zero through 1977 and one thereafter, interactively with the seasonal dummies.

Figure 1 contains a block of four graphs, labelled as Figures 1a–d, which plot (*hi*, *rp*), (*hi*, –*z*), (*hi*, –*s*) and (*CU**, *hi**) over the sample; *CU** is *CU* adjusted for deterministic seasonality, and *hi** is *hi* adjusted for deterministic seasonality and trend. To further highlight correlations in the data, *rp*, –*z* and –*s* are adjusted to match the mean of *hi*, and *hi** is adjusted to match the mean and range of *CU**. Figure 1a shows that *hi* and *rp* have clear downward trends. The output-imports ratio decreases by 70% and *RP* by 34% from 1968 to 1994, suggesting a long-run substitution elasticity above two if

**Figure 1. Time series for pairs of variables**



$\mu_2 = \mu_3 = 0$. However, Figure 1c, displaying positive and strong correlation between $-s$ and the trend in *hi*, indicates that increased specialisation accounts for much of the fall in *hi*. Likewise, but not shown, *H* over manufacturing production decreases from 0.84 in 1968 to 0.69 in 1994. The development of $-z$ also matches the decline in *hi* through 1987; but, while *hi* continues to fall thereafter, $-z$ increases from 1987 to 1994 (see Figure 1b).

Consistent with the $\mu_4 < 0$ hypothesis, Figure 1d shows that $CU^*$ and $hi^*$ are negatively correlated. Specifically, the 1984–1987 boom shows up in high operating rates and low values of $hi^*$; and the subsequent slump is accompanied by high $hi^*$ values and low rates of utilisation. The figure also suggests that *CU* is I(1) over the sample, but this hypothesis is rejected by augmented Dickey Fuller tests. Furthermore, it may be argued that *CU* is I(0) by construction.

## 4. Cointegration Analysis

The analysis commenced from a fifth-order vector autoregression (VAR) in *hi*, *rp*, *z*, *s*, and *CU* which included an unrestricted intercept and the aforementioned dummies. That VAR had statistically acceptable equations for $hi_t$, $rp_t$, $z_t$ and $s_t$, but the $CU_t$ equation suffered from residual autocorrelation and coefficient instability. Recognising this problem and treating *CU* as I(0), we re-specified the VAR by conditioning on the lagged utilisation terms. The modified system was then (validly) simplified to a conditional VAR with three lags on $CU_t$ and the modelled variables. This system is used below, as further simplification yielded a $z_t$ equation with autocorrelated residuals.

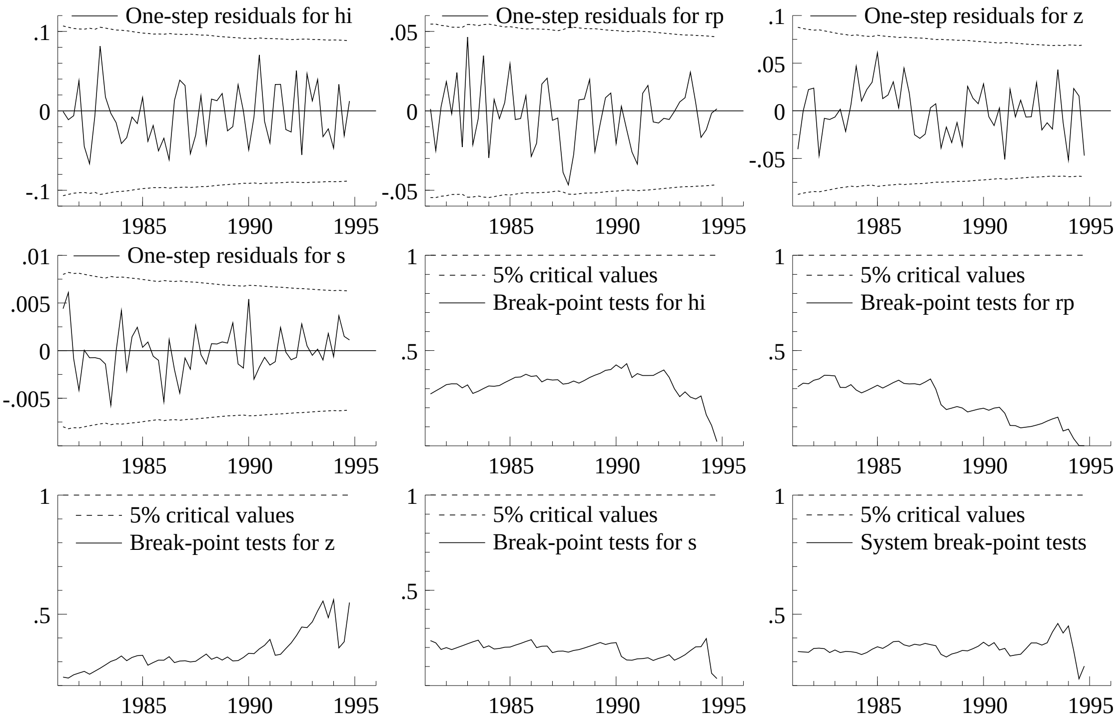
Univariate and multivariate diagnostic tests for the third-order VAR are reported in Table 1. The diagnostics are satisfactory, even though an outlier induces significant non-normality in the equation for $s_t$. That is, the non-normality is not critical for the Johansen analysis, particularly if (as we expect) $s_t$ is weakly exogenous for the cointegration parameters; see Johansen and Juselius (1992, p. 219), Cheung and Lai (1993, Section 4) and Johansen (1995, p. 29). Next, Figure 2 plots the one-step residuals with 0±2 equation standard errors and sequences of break-point Chow tests (scaled by their 5% critical values). The constancy statistics are only descriptive, since the system is in I(1) space, but they indicate that the coefficients are stable.[12] In conclusion, the system seems to be a valid baseline for further analysis.

**Table 1. Diagnostic Tests**

| Statistic | *hi* | *rp* | *z* | *s* | *VAR* |
|---|---|---|---|---|---|
| $AR_{1\text{-}5}$ : $F(5, 80)$ | 0.53 | 1.89 | 0.86 | 1.56 | |
| $NORM$ : $\chi^2(2)$ | 4.87 | 0.29 | 1.15 | 7.46* | |
| $ARCH_{1\text{-}4}$ : $F(4, 77)$ | 0.64 | 0.20 | 0.43 | 0.22 | |
| $HET$ : $F(24, 60)$ | 0.89 | 0.72 | 1.01 | 0.54 | |
| $AR^{M}_{1\text{-}5}$ : $F(80, 247)$ | | | | | 1.05 |
| $NORM^{M}$ : $\chi^2(8)$ | | | | | 11.10 |
| $HET^{M}$ : $F(240, 497)$ | | | | | 0.79 |

Notes: $F(\cdot,\cdot)$ and $\chi^2(\cdot)$ denote approximate null distributions, and the degrees of freedom are in parentheses. $AR_{1\text{-}5}$ is a Lagrange multiplier test for fifth-order residual autocorrelation (see Harvey (1981)); *NORM* is the normality test suggested by Doornik and Hansen (1994); $ARCH_{1\text{-}4}$ is a Lagrange multiplier test for fourth-order ARCH residuals due to Engle (1982); and *HET* checks whether the squared residuals depend on the regressors and their squares (cf. White (1980)); corresponding multivariate (system) tests are denoted by $^{M}$ (see Doornik and Hendry (1997)). Following Gerrard and Godfrey (1998, Section 3), the *HET* statistics are calculated for the equilibrium correction form of the system. Here and below, * and ** denote significance at 5% and 1% respecively. The quantiles used here are not adjusted to match the I(1)-ness of the series (cf. Doornik and Hendry (1997, p. 37)).

**Figure 2. Recursive test statistics for the system**

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[12] Another caveat is that the coefficients of the unrestricted variables (the intercept, the dummies and the *CU*-terms) are fixed at their full-sample values (cf. Doornik and Hendry (1997, p. 271)). The impression from Figure 2 is not changed if all the coefficients are allowed to vary.

**Table 2. Cointegration Test Statistics**

| Hypothesis | Eigenvalues: 0.278, 0.165, 0.079, 0,011 | | | | | |
|---|---|---|---|---|---|---|
| | $\lambda_{max}$ | $\lambda^a_{max}$ | 95% quantile | $\lambda_{trace}$ | $\lambda^a_{trace}$ | 95% quantile |
| $r = 0$ | 34.9** | 31.0* | 27.1 | 64.2** | 57.0** | 47.2 |
| $r \leq 1$ | 19.3 | 17.1 | 21.0 | 29.3 | 26.0 | 29.7 |
| $r \leq 2$ | 8.8 | 7.8 | 14.1 | 10.1 | 8.9 | 15.4 |
| $r \leq 3$ | 1.2 | 1.1 | 3.8 | 1.2 | 1.1 | 3.8 |

Notes: The statistics $\lambda_{max}$ and $\lambda_{trace}$ are the maximum eigenvalue and trace statistics for determining the cointegration rank, $r$; $\lambda^a_{max}$ and $\lambda^a_{trace}$ are parallel statistics with a small-sample correction; cf. Reimers (1992) and Cheung and Lai (1993, Section 3). The quantiles are asymptotic, and are from Osterwald-Lenum (1992, Table 1).

Table 2 contains results from applying Johansen's (1988, 1991) method to the preferred system. The maximum eigenvalue and trace statistics reject the null of no cointegration at the 1% level; and parallel statistics with a small-sample correction reject at 5% and 1%.[13] The null of at most one I(0) relation is not rejected at 5%, so we proceed on the assumption of a single cointegrating vector among ($hi$, $rp$, z, $s$).[14] Normalised on $hi$, the unrestricted estimate of the vector is given by:

(2) $hi = intercept + 1.562\, rp - 0.077\, z - 0.613\, s,$

and the corresponding vector of feedback coefficients, denoted $\hat{\alpha}'$, is:

(3) $\hat{\alpha}' = (-0.390, 0.037, -0.067, -0.003).$

The estimates in (2) and (3) suggest that the cointegrating relation is an output-imports model consistent with arguments above. First, disequilibrium in the relation appears to feed back strongly onto $hi$ and weakly, if at all, onto $rp$, $z$ and $s$. Second, the coefficients of $rp$ and $s$ in (2) have their

[13] Strictly speaking, the quantiles in Table 2 are invalid, even asymptotically. That is, the tests are not asymptotically similar when applied to (standard) systems with I(0), unrestricted conditioning variables; see Rahbek and Mosconi (1998). This difficulty is resolved by Rahbek and Mosconi's (1998) approach, but using that approach is beyond the scope of our study. The *CU* terms were significant, so exclusion of the terms could lead to inefficient estimates of the cointegration parameters.

[14] The second eigenvector may also be I(0), however. Allowing for two cointegrating vectors, we identify a vector almost identical to the restricted vector (4) below and a vector including $rp$, $z$ and $s$. In accordance with the results in Table 3, the first vector appears to enter the $hi_t$ equation only (but it might also enter the equation for $z_t$), whereas the second vector is significant only in the $z_t$ equation. Imposing the corresponding (six) exogeneity restrictions, we obtain an estimate of the first vector which is close to the estimate in (4). In short, the choice between one or two I(0) relations is not crucial for the below analysis.

expected signs, and they seem reasonable from the discussion of Figure 1. Lastly, the expenditure elasticity is numerically close to zero, indicating that $z$ may be omitted from (2).

This interpretation is supported by the statistics for testing weak exogeneity and long-run exclusion in Table 3 – each of which is asymptotically distributed as $\chi^2$ with degrees of freedom in parentheses. The exogeneity tests show that $rp_t$, $z_t$ and $s_t$ can be considered weakly exogenous for the cointegration parameters, but long-run weak exogeneity of $hi_t$ is clearly rejected. Moreover, each of $hi$, $rp$ and $s$ is statistically significant in the cointegrating vector – at least conditionally on ($rp_t$, $z_t$, $s_t$) being weakly exogenous for that vector – whereas $z$ is far from being significant. Imposing a zero long-run coefficient on $z$ and weak exogeneity of ($rp_t$, $z_t$, $s_t$), we obtain $\chi^2(4) = 1.52$ (with a $p$-value of 0.82) and the following estimate of the cointegrating vector:

$$(4) \quad hi = intercept + \underset{(0.265)}{1.475}\, rp - \underset{(0.139)}{0.717}\, s,$$

where estimated standard errors are in parentheses. The coefficients in (4) have absolute $t$-values above 5 and; and they are close to the comparable estimates in (2), supporting the validity of the reduction. The price elasticity implies that 100·*IMP* decreases by 0.37 percentage points if *RP* increases permanently by 1% in 1994; *I* declines by 0.75%. These estimates are in line with long-run price coefficients in other studies of manufacturing imports or import shares; see Menon (1995), the UK studies cited in footnote 1 and Haas and Turner's (1990, Table 10) models for 14 OECD countries (including Norway). As regards the specialisation effects, (4) implies that *I* and 100·*IMP* decrease by 0.36% and 0.18 points if *S* rises permanently by 1% in 1994 (*ceteris paribus*); *S* grew by 4% that year. Recent UK studies, in contrast, obtain long-run specialisation elasticities in the range 0.8-1.6 for manufactured imports (see Anderton *et al*. (1992), Sedgley and Smith (1994) and Temple and Urga (1997)).

As Figure 3 shows, the recursively estimated parameters of $rp$ and $s$ in (4) are constant.[15] Their approximate 95% confidence intervals are wide, however, owing to the high correlation between $rp$ and $s$ (cf. Figure 1). This correlation probably reflects that $rp$ and $s$ are driven by some of the same factors. Nonetheless, it seems useful to compute their separate contributions to the 1968–1994 decline in $hi$.

**Table 3. Tests for Long-Run Weak Exogeneity and Long-Run Exclusion**

[15] Following Hansen and Johansen (1993), these estimates were obtained after fixing the short-run parameters and the coefficients of unrestricted variables at their full-sample values.

| Variables | Long-run weak exogeneity<br>Likelihood ratio (df) | <br>*p*-value |
|---|---|---|
| *hi* | $\chi^2(1) = 12.06^{**}$ | 0.00 |
| *rp* | $\chi^2(1) = 0.65$ | 0.42 |
| *z* | $\chi^2(1) = 0.47$ | 0.49 |
| *s* | $\chi^2(1) = 0.19$ | 0.66 |
| (*rp*, *z*, *s*) | $\chi^2(3) = 1.42$ | 0.70 |

| Variables | Long-run exclusion<br>Likelihood ratio (df) | <br>*p*-value |
|---|---|---|
| *hi* | $\chi^2(1) = 12.69^{**}$ | 0.00 |
| *rp* | $\chi^2(1) = 7.40^{**}$ | 0.01 |
| *z* | $\chi^2(1) = 0.05$ | 0.82 |
| *s* | $\chi^2(1) = 2.55$ | 0.11 |
| (*z*, *s*) | $\chi^2(2) = 10.14^{**}$ | 0.01 |

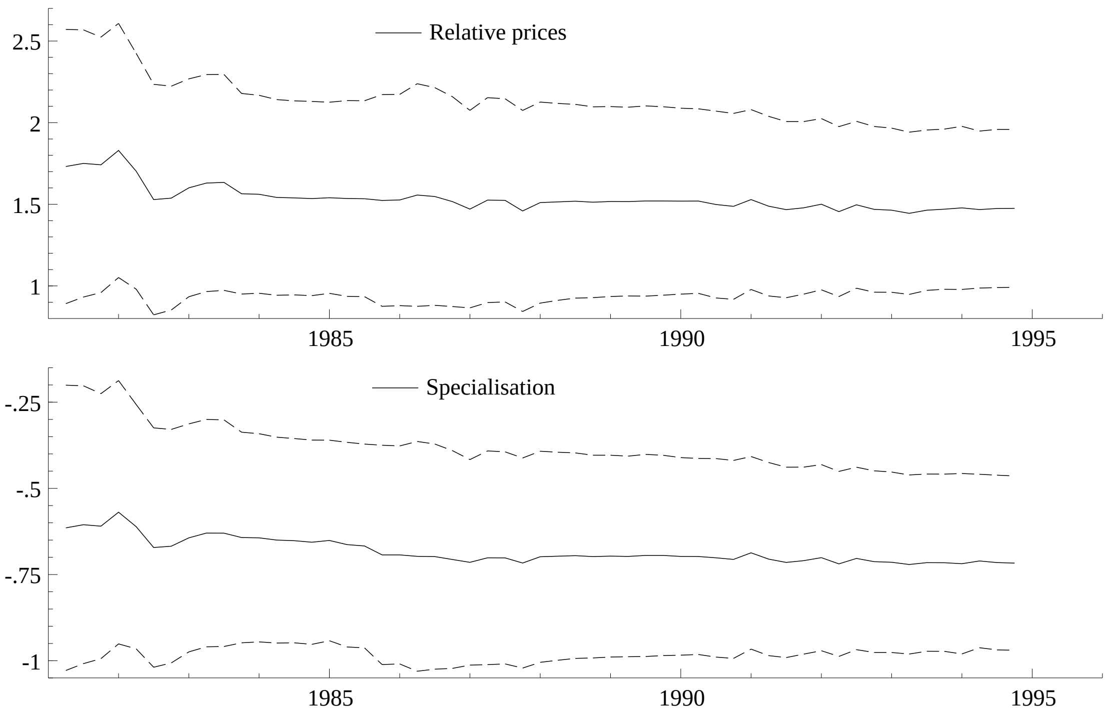
| Variable | Long-run exclusion conditional on (*rp*, *z*, *s*) being weakly exogenous for the cointegration parameters<br>Likelihood ratio (df) | <br>*p*-value |
|---|---|---|
| *hi* | $\chi^2(1) = 19.98^{**}$ | 0.00 |
| *rp* | $\chi^2(1) = 10.33^{**}$ | 0.00 |
| *z* | $\chi^2(1) = 0.10$ | 0.75 |
| *s* | $\chi^2(1) = 7.79^{**}$ | 0.01 |

Notes: The statistics are calculated under the assumption that $r = 1$, and are asymptotically distributed as $\chi^2$(df) under the null. The conditional exclusion statistics parallel statistics in Johansen and Juselius (1990, Section 6.1).

On the basis of (4) and the average quarterly changes in *rp* and *s*, 50-55% (45-50%) of the decline is explained by the reduction in *rp* (growth in *s*).

The feedback coefficient corresponding to (4) is –0.442, meaning that 44% of a disequilibrium in (4) is corrected within one quarter. Although informative about disequilibrium correction, the feedback coefficient only gives limited information about the short-run adjustment of *hi*. That adjustment is of interest from a policy perspective, so it is investigated below.

**Figure 3. Equation (4): recursive estimates (—) with ± 2 estimated standard errors (---)**




# 5. An Equilibrium Correction Model of the Output-Imports Ratio

We now focus on (a) the direct long-run effects on *hi* of changes in capacity utilisation, and (b) the dynamic adjustment of *hi* to changes in expenditure, utilisation and (relative) prices. In light of the results in Section 4, (a) and (b) are examined with a single equation model of *hi* which includes disequilibrium in (4) as an equilibrium correction term. Using the same lag length as in the initial VAR of Section 4, and allowing for separate short-run effects of *pi* and *ph*, the single equation analysis commenced from the following specification:

$$(5)\quad \Delta hi_t = \beta_0 + \sum_{i=0}^{4} \beta_{1i} \Delta pi_{t-i} + \sum_{i=0}^{4} \beta_{2i} \Delta ph_{t-i} + \sum_{i=0}^{4} \beta_{3i} \Delta z_{t-i} + \sum_{i=0}^{4} \beta_{4i} \Delta CU_{t-i}$$
$$+ \sum_{i=1}^{4} \beta_{5i} \Delta hi_{t-i} + \beta_6 CU_{t-1} + \gamma (hi_{t-1} - 1.475 rp_{t-1} + 0.717 s_t)$$
$$+ \sum_{i=1}^{3} (\beta_{8i} Q_{it} + \beta_{9i} D_t Q_{it}) + \varepsilon_t ,$$

where $\varepsilon_t$ is the error term, $\Delta$ is the difference operator and ($hi_{t-1} - 1.475\ rp_{t-1} + 0.717\ s_t$) is the equilibrium correction mechanism. Equation (5) was simplified to the model (6) using a general-to-specific procedure.[16] The simplification was conducted with instruments for $\Delta z_t$ and $\Delta CU_t$, to take account of the simultaneity between these variables and $\Delta hi_t$.[17] Both $\Delta z_t$ and $\Delta CU_t$ are omitted from (6), however, since they were statistically insignificant. So were their lags, the lags on ($\Delta pi_t$, $\Delta ph_t$, $\Delta hi_t$) and tests of short-run price homogeneity. Consequently, the preferred equation only includes three economic variables: $\Delta rp_t$, $CU_{t-1}$ and the equilibrium correction term. Conditional on the weak exogeneity of $\Delta rp_t$ for the cointegration parameters, a Wu-Hausman test indicated that this variable is weakly exogenous for all the parameters in (6).[18] It is thus estimated by least squares.

$$
\begin{aligned}
(6)\quad \Delta hi_t \;=\;& \underset{(0.083)}{0.547} + \underset{(0.161)}{0.833}\,\Delta rp_t - \underset{(0.088)}{0.551}\,CU_{t-1} \\
& - \underset{(0.060)}{0.436}\,(hi_{t-1} - 1.475\,rp_{t-1} + 0.717\,s_t) \\
& - \underset{(0.008)}{0.058}\,(Q_2 + Q_3)_t + \underset{(0.010)}{0.031}\,(Q_1 - D{\cdot}Q_1 - D{\cdot}Q_2)_t
\end{aligned}
$$

$T = 107$ [1968(2) – 1994(4)] $\quad R^2 = 0.625 \quad \sigma = 3.967\% \quad DW = 2.29$

$AR_{1\text{-}5} : F(5, 96) = 1.27 \quad NORM : \chi^2(2) = 0.33 \quad ARCH_{1\text{-}4} : F(4, 93) = 0.56$

$HET : F(8, 92) = 0.70 \quad HET_G : F(18, 82) = 0.83 \quad RESET : F(1, 100) = 4.80^*$.

Estimated standard errors are in parentheses, and $T$, $R^2$, $\sigma$ and $DW$ respectively are the number of observations, the squared multiple correlation coefficient, the residual standard deviation and the Durbin-Watson statistic. In addition to the tests defined in Table 1, (6) reports $HET_G$, which checks whether the squared residuals depend on the levels, squares and cross-products of the regressors (see White (1980)), and the Ramsey (1969) test for functional form mis-specification, *RESET*. The latter has a *p*-value of 0.03, but the other tests are insignificant at conventional levels. Hence the model appears reasonably well specified on the diagnostics.

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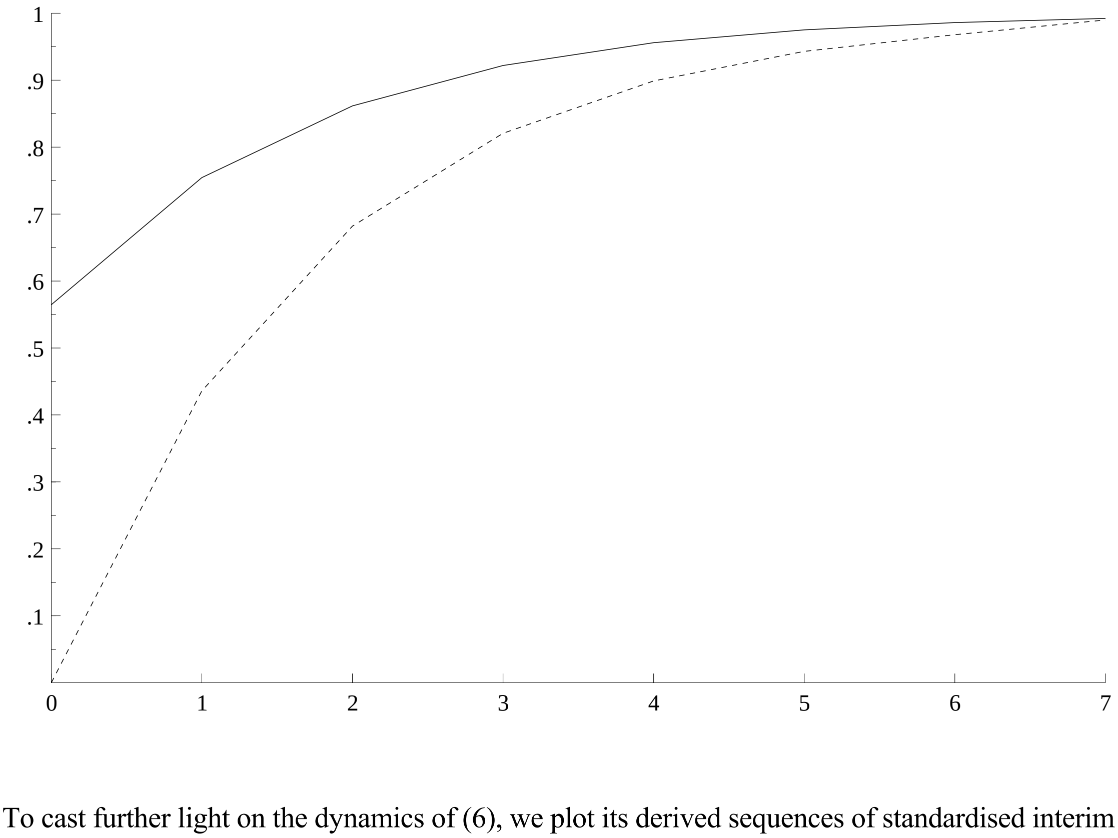
[16] Since many hypotheses were tested during the simplification, we only retained economic variables (or sets of variables) that were significant at 0.1% (cf. Hendry (1995b, Section 13.10.5)).

[17] This simultaneity does not preclude $z_t$ from being weakly exogenous for the cointegration parameters; see Urbain (1992b, 1995) and Hendry (1995a).

[18] The test was based on reduced forms for $\Delta pi_t$ and $\Delta ph_t$ as estimated $\Delta rp_t$ equations were hard to interpret. The model of $\Delta pi_t$ included the following economic variables, with *t*-values in parentheses: $\Delta_3 pf_t$ (5.6), $\Delta_3 e_t$ (5.1), $u_{t-1}$ (–7.9), ($\Delta pi_{t-1} + \Delta pi_{t-4}$) (–4.3) and ($pi_{t-1} - 0.483\, pf_{t-2} - 0.296\, e_{t-1} - 0.403\, ph_{t-1}$) (–7.0); the $\Delta ph_t$ model contained $\Delta pi_{t-1}$ (2.0), $\Delta_2 uc_{t-2}$ (1.9), $\Delta ph_{t-2}$ (–2.5), $CU_{t-1}$ (2.6) and $(ph - 0.159\, pi - 0.841\, uc)_{t-1}$ (–2.9); see the Appendix and Section 6 for details. Adding the derived $\Delta rp_t$ estimate to (6) yielded a *t*-value of 0.67. Furthermore, ($hi_{t-1} - 1.475\, rp_{t-1} + 0.717\, s_t$) was insignificant when added

The economic variables in (6) are highly significant, with absolute *t*-values exceeding 5. Relative prices *rp* has a positive impact coefficient which is numerically large and more than one half of the price elasticity in (4). Lagged utilisation also obtains an economically important coefficient, and the long-run semi-elasticity of *CU* is –1.264. Combined with Figure 1d, this estimate implies that variation in *CU* induced marked, direct *hi* changes over 1968–1994. It also implies that 100·*IMP* increases by 0.32 percentage points if 100·*CU* increases permanently by 1 percentage point in 1994 (other things being equal); *I* rises by 0.64%; and *H* decreases by 0.62%. The equilibrium correction coefficient is close to the restricted feedback elasticity in Section 4, and its *t*-value (–7.3) reinforces our I(0) interpretation of ($hi$ – 1.475 $rp$ + 0.717 $s$). Finally, the significant dummies are merged into two terms in order to simplify the model and to facilitate recursive estimation.

**Figure 4. Equation (6): standardised interim multipliers for $rp_t$ (—) and $CU_t$ (---)**




To cast further light on the dynamics of (6), we plot its derived sequences of standardised interim multipliers for $rp_t$ and $CU_t$. Figure 4 shows that *hi* adjusts rapidly to changes in *rp*: 75% of the total adjustment is completed within two quarters, and 92% of the adjustment has come through after one year.

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to the marginal models, confirming that $rp_t$ is weakly exogenous for the cointegration parameters in (4)-(6). Long-run

These multipliers indicate that purchasers are reasonably well informed about $rp_t$; do not face long contracts and delivery lags; and only to a limited extent have habits and/or adjustment costs which slow down the adjustment. The response to *CU* changes is more gradual, but 90% of the long-run response is completed after five quarters. Since *CU* proxies for factors that, together with *RP*, may be viewed as elements of *effective* relative prices (see Gregory (1971)), one would expect *hi* to respond symmetrically to *observed* changes in those factors and *rp*. Apparently, but not surprisingly, agents have better information on *rp* than on other components of effective prices.

## 6. Testing the Lucas Critique

This section evaluates whether or not equation (6) suffers from the Lucas critique. Interpreted broadly, Lucas's critique may apply to (6) at different levels (cf. Favero and Hendry (1992, pp. 267-68)). Specifically, since it is specified and estimated as a contingent plan equation, (6) will "break down" if agents use expectations models of *rp* and those models change sufficiently; see Hendry (1988) and Favero and Hendry (1992). Furthermore, even if expectations formation is not at issue, the parameters in (6) need not be invariant to all conceivable interventions. In particular, one may suspect agents to react more quickly to large *rp* changes than to small changes (cf. Orcutt (1950, p. 126), Liu (1954, p. 437) and Goldstein and Khan (1976)). The argument is that (i) large changes often are viewed as more permanent than small ones; and (ii) shifts which appear long-lived are, in the presence of adjustment costs, likely to yield faster response than changes which might be reversed rapidly.

Following Hendry (1988) and Engle and Hendry (1993), a way to proceed would be to test the invariance of equation (6)'s parameters to changes in the marginal process for $rp_t$. That process appears reasonably constant over the sample, however, so other approaches must be used.[19] We first apply variable-addition tests to examine the role of expectations; cf. Hendry (1988, p. 135), Hendry and Neale (1988, p. 813) and, in particular, Engle and Hendry (1993, pp. 127-31). Then tests and measures of parameter constancy are used to check if (6) is autonomous to unspecified interventions (i.e., the interventions which occurred over 1968–1994). Lastly, the hypothesis of an invariant adjustment to *rp* is tested within Liu's (1954) and Goldstein and Khan's (1976) framework.

It is first noted that the Wu-Hausman test in Section 5 is informative about expectations formation. In fact, it refutes the claim that agents act on expectations models of $\Delta rp_t$. Its alternative hypothesis does not allow for model-based expectations of $\Delta rp_{t+1}$, however, a possibility which should be considered.

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homogeneity was rejected for the $\Delta pi_t$ equation.

We do this by testing $\Delta rp_{t+1}$ (properly instrumented) or determinants of $\Delta rp_{t+1}$ for inclusion in (6): significance of the added variables would be evidence in support of Lucas's critique. These tests cannot be based on the above models of $\Delta pi_t$ and $\Delta ph_t$ (cf. footnote 18), as their $\Delta rp_{t+1}$ predictions involve leaded variables and other terms that would be endogenous in (6). The next step of the analysis, therefore, is to develop new price equations.

Starting from general specifications and drawing on results in Naug and Nymoen (1996), we obtained the following models of $\Delta pi_t$ and $\Delta ph_t$ (seasonal dummies are omitted to save space):

$$\text{(7)}\quad \Delta pi_t = -\underset{(0.013)}{0.105} + \underset{(0.057)}{0.144}\,\Delta_2 pf_{t-1} + \underset{(0.066)}{0.242}\,\Delta_2 e_{t-1}$$

$$-\underset{(0.004)}{0.036}\,u_{t-2} - \underset{(0.054)}{0.226}\,(\Delta pi_{t-1} + \Delta pi_{t-4})$$

$$-\underset{(0.058)}{0.447}\,(pi - 0.483\,pf - 0.296\,e - 0.403\,ph)_{t-1}$$

$T = 107$ [1968(2) – 1994(4)] $\quad R^2 = 0.564 \quad \sigma = 1.330\% \quad DW = 1.97$

$AR_{1\text{-}5} : F(5, 94) = 1.48 \quad NORM : \chi^2(2) = 5.37 \quad ARCH_{1\text{-}4} : F(4, 91) = 0.58$

$HET : F(14, 84) = 0.38 \quad HET_G : F(34, 64) = 0.58 \quad RESET : F(1, 98) = 0.21$

and

$$\text{(8)}\quad \Delta ph_t = -\underset{(0.024)}{0.038} + \underset{(0.072)}{0.198}\,\Delta pi_{t-1} + \underset{(0.087)}{0.215}\,\Delta uc_{t-3}$$

$$+\underset{(0.027)}{0.064}\,CU_{t-2} - \underset{(0.066)}{0.160}\,\Delta_2 ph_{t-1}$$

$$-\underset{(0.048)}{0.134}\,(ph - 0.159\,pi - 0.841\,uc)_{t-2}$$

$T = 107$ [1968(2) – 1994(4)] $\quad R^2 = 0.415 \quad \sigma = 1.288\% \quad DW = 1.92$

$AR_{1\text{-}5} : F(5, 94) = 0.48 \quad NORM : \chi^2(2) = 0.35 \quad ARCH_{1\text{-}4} : F(4, 91) = 2.27$

$HET : F(12, 86) = 0.84 \quad HET_G : F(33, 65) = 1.17 \quad RESET : F(1, 98) = 3.14.$

Consonant with Naug and Nymoen (1996, eq. 14), the $\Delta pi_t$ model includes effects of *ph*, foreign export prices (*pf*), the exchange rate (*e*) and the domestic rate of unemployment (*u*), which is an easily

[19] The $rp_t$ equation of the VAR and simplifications of that equation are empirically constant without intervention dummies (cf. Figure 2). So are the marginal models of $\Delta pi_t$ and $\Delta ph_t$ reported below and in footnote 18.

observable proxy for demand pressure.[20] In (7), *pi* adjusts considerably to changes in *ph* and *u*; and shifts in *px* and *e* are (other things being equal) not fully passed through to *pi*. The model for $\Delta ph_t$ contains effects of *pi*, *CU* and domestic costs (*uc*). Interestingly, the long-run *CU* coefficient (0.48) is slightly smaller in absolute value than our long-run semi-elasticity of *H* with respect *CU* (–0.62).

Equations (7) and (8) include six regressors which were not used in the design of (6). Adding their one-period leads to (6) yields $F(6, 95) = 1.06$ and a *p*-value of 0.39; none of the added terms is significant.[21] Similarly, the derived $\Delta rp_{t+1}$ estimate obtains a *t*-value of 0.99 ($p = 0.33$) when included in (6). In short, we do not find significant evidence in favour of the expectations hypothesis.

Nevertheless, (6) has a forward-looking interpretation if agents employ $\Delta rp_t$ as a data-based predictor of $\Delta rp_{t+1}$; see Hendry and Ericsson (1991, Section 4.3) and Favero and Hendry (1992, Section 9). Such behaviour could be rational if the expected gain from using a model-based predictor (rather than $\Delta rp_t$) is smaller than the expected information costs of obtaining it. Predicting $\Delta rp_{t+1}$ by $\Delta rp_t$ works poorly within the sample, however: the standard deviation of the prediction error is 4.1%, while the unconditional standard deviation of $\Delta rp_t$ is 2.5%. In other words, agents appear to act contingently on the observed values of $\Delta rp_t$.

The long-run coefficients of *rp* and *s* are already found to be empirically constant. We now investigate the constancy or otherwise of equation (6), taking the estimated cointegration parameters in (4) as known. Figure 5 shows the one-step residuals with $0 \pm 2\sigma_t$ and the sequence of break-point Chow statistics (scaled by their 5% critical values). The residual standard error is stable, and all of the Chow statistics are insignificant at the 5% level. Further, none of the Hansen (1992) tests in Table 4 is significant at 10%. Figure 6 plots the recursive estimates with $\pm 2$ standard errors for the coefficients on the economic variables. It is seen that the estimates are stable and statistically significant over 1973–1994; the other coefficients in (6) are also reasonably constant, albeit less so than those plotted in Figure 6.

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[20] The long-run estimates in (7) differ markedly from those in Naug and Nymoen (1996), however. The cause of the difference is that Naug and Nymoen (1996) (a) model a broader aggregate than *pi*, and (b) use domestic wage costs instead of domestic prices. The hypothesis of long-run homogeneity is rejected for (7).

[21] The *p*-value of the *F* test becomes 0.26 if the (six) seasonal dummies are included unrestrictedly in (6).

**Figure 5. Recursive test statistics for equation (6)**

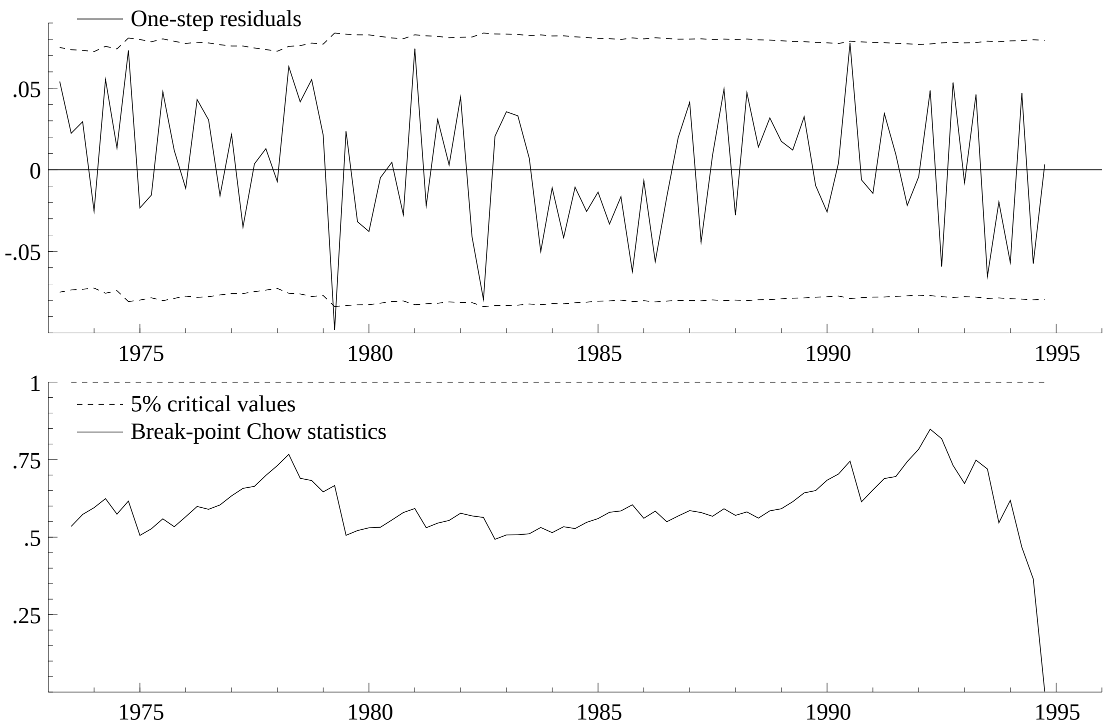


**Figure 6. Equation (6): recursive estimates (—) of the coefficients on the economic variables, with ± 2 estimated standard errors (---)**

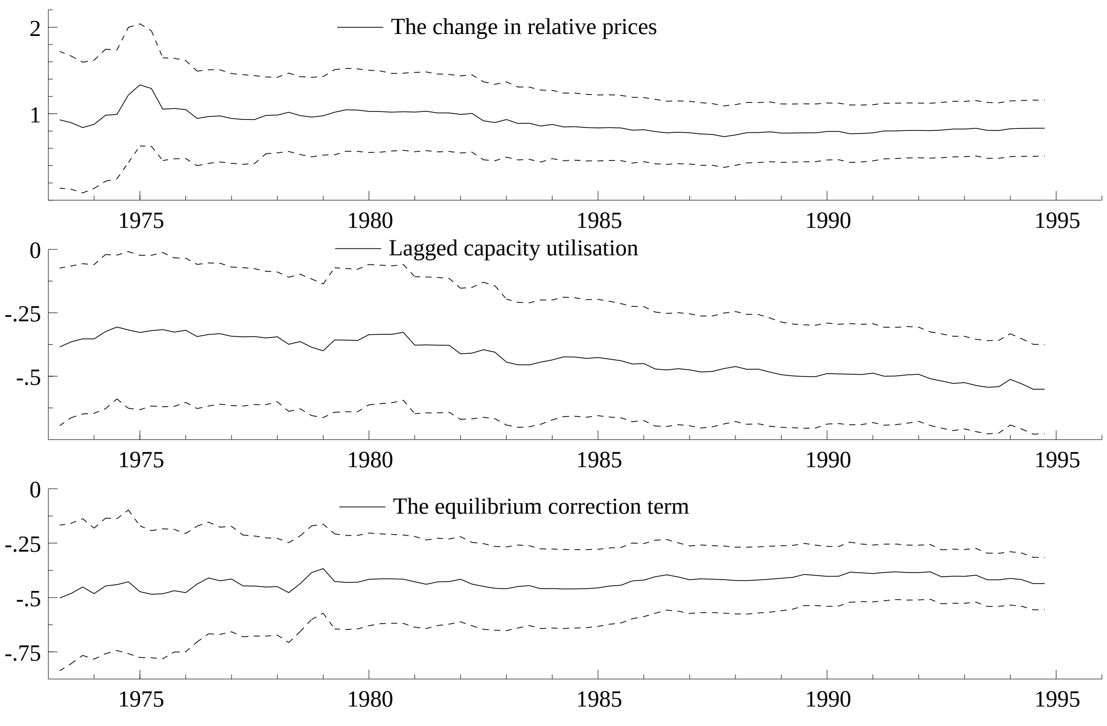

**Table 4. Hansen (1992) Tests for Parameter Instability with an Unknown Break Point**

| Regressor/parameter(s) in equation (6) | Test-value |
|---|---|
| Constant term | 0.09 |
| $\Delta rp_t$ | 0.07 |
| $CU_{t-1}$ | 0.10 |
| $(hi_{t-1} - 1.475\, rp_{t-1} + 0.717\, s_t)$ | 0.07 |
| $(Q_2 + Q_3)_t$ | 0.08 |
| $(Q_1 - D{\cdot}Q_1 - D{\cdot}Q_2)_t$ | 0.28 |
| Residual variance | 0.04 |
| All parameters jointly | 1.30 |

Notes: The first seven statistics test the stability of each parameter individually, and the last statistic tests the stability of all the parameters jointly. An individual stability test rejects the null of parameter constancy at 5 per cent (10 per cent) if the test-value exceeds 0.47 (0.353). The 5% and 10% critical values for the joint stability test respectively are 1.90 and 1.69. The critical values are asymptotic, and are from Hansen (1992, Table 1).

We finally test whether agents adjust faster to large changes in *rp* than to small changes. The proposed alternative to invariance is that $\beta_{1t} = \pi_0 + \pi_1\, |\Delta rp_t|$ and $\gamma_t = \pi_2 + \pi_3\, |\Delta rp_{t-1}|$, where $\pi_1 > 0$, $\pi_3 < 0$, $\beta_{1t}$ is the parameter on $\Delta rp_t$ and $\gamma_t$ is the parameter on $(hi_{t-1} - 1.475\, rp_{t-1} + 0.717\, s_t)$. Hence the approach is to test $|\Delta rp_t|{\cdot}\Delta rp_t$ and $|\Delta rp_{t-1}|{\cdot}(hi_{t-1} - 1.475\, rp_{t-1} + 0.717\, s_t)$ for significance in (6). Again, the added variables are individually and jointly insignificant, with joint test yielding $F(2, 99) = 1.19$ ($p = 0.31$). We are thus unable to reject invariance in favour of the specified alternative. Goldstein and Khan (1976) obtain a similar result for import demand in 12 OECD countries (including Norway).

# 7. Concluding Remarks

This paper has modelled domestic output over imports in Norway's expenditure on manufactures. We first obtained a cointegrating vector for the output-imports ratio with significant effects of relative prices and a proxy for international specialisation; domestic activity, by contrast, was insignificant. A restricted vector (without activity) yielded a long-run substitution elasticity of 1.48 and a specialisation elasticity of –0.71. Based on these estimates, we calculated that increasing specialisation (*ceteris paribus*) explains 45-50%, and decreasing import prices to domestic prices 50-55%, of the 1968–1994 decline in the output-imports ratio. The absence of long-run activity effects is in line with the few import studies which have allowed for effects of activity, specialisation and capacity utilisation. Other import studies typically find that import shares increase with domestic demand in the long run. Thus, demand changes appear to have stronger long-run effects on manufacturing output – and weaker effect on manufactured imports – than suggested by standard import models. However, unless the trend towards increased specialisation halts, the results also mean that domestic manufacturing becomes less and less sensitive to domestic demand conditions. This tendency is reinforced by the declining share of manufactures in domestic expenditure.

We next developed a conditional equilibrium correction model of the output-imports ratio, using the restricted cointegrating vector in the equilibrium correction term. In addition to strong disequilibrium correction effects, the model included important short-run influences of relative prices and marked negative effects of domestic capacity utilisation. The estimates imply that the output-imports ratio adjusts rapidly to changes in relative prices and capacity utilisation.

We also examined the model in light of Lucas's (1976) critique. Stated briefly, the critique was not supported by the data. In particular, variable-addition tests refuted the claim that agents act on expectations models of relative prices. An explanation of this result is that it is costly/difficult to derive adequate models of Norwegian import prices: as shown here and in Naug and Nymoen (1996), such models are more complex than suggested by conventional theory. The output-imports model was also found to be constant over time and invariant to large changes in relative prices. That said, it might still break down in the face of future interventions.

Bearing this caveat in mind, the estimated price effects have implications for policy. First and foremost, manufacturing output is sensitive to policies which change relative prices. On this basis, and assuming output variability to be undesirable, one may argue against policy regimes that allow the currency to fluctuate. Such an argument is weakened by our price models, however, which imply that exchange rate changes are far from yielding proportional changes in relative prices. A similar reasoning applies to the utilisation effects. From the output-imports model, changes in capacity utilisation have strong, direct influences on manufacturing output and the import share. Hence it would be unsurprising if utilisation changes had weak effects on prices. This only seems to be true for *relative* prices, however; both domestic prices and import prices appear to vary pro-cyclically with demand pressure in Norway.

# Data Definitions

- $HI$ = The output-imports ratio at fixed 1991 basic prices. It is constructed as $HI = H/I = (X - A)/I$, where $X$, $A$ and $I$ respectively are gross production, exports and imports (including tariffs) of manufactures. The series exclude ships, metals, oil platforms, basic chemicals, pulp and paper, refined oil products and non-competing imports. Source: Statistics Norway (Quarterly national accounts, QNA).
- $PI$ = Implicit price deflator of $I$ (1991 = 1). Source: Statistics Norway (QNA).
- $PH$ = Implicit price deflator of $H$ (1991 = 1). Source: Statistics Norway (QNA).
- $RP = PI/PH$.
- $Z = H+I$.
- $S$ = Eight-quarter moving average of manufacturing exports to manufacturing production – both measured in volume terms – in OECD countries (excluding Norway and recent OECD entrants). The corresponding non-smoothed variable equals 1 (on average) in 1990. The countries in the index are weighted according to their importance in OECD exports of goods. Source: OECD (Main Economic Indicators).
- $CU$ = Capacity utilisation in the production of $X$ (see Section 3). Source: Statistics Norway.
- $Q_{it}$ = Centred dummy for quarter i, equals 0.75 in quarter i, –0.25 otherwise. i = 1,2,3.
- $D_t$ = Step dummy for a structural break in the seasonal pattern of the QNA series, equals zero through 1977, one thereafter.
- $CU^*$ = The residual plus the estimated intercept from regressing $CU$ on an intercept and the variables ($Q_1$, $Q_2$, $Q_3$, $Q_1 \cdot D$, $Q_2 \cdot D$, $Q_3 \cdot D$).
- $hi^*$ = The residual from regressing $hi$ on an intercept, a linear trend and ($Q_1$, $Q_2$, $Q_3$, $Q_1 \cdot D$, $Q_2 \cdot D$, $Q_3 \cdot D$). The mean of the series is adjusted to match that of $CU$.
- $PF$ = Index of foreign export prices (1991 = 1). The index covers merchandise exports of the 14 countries with largest weights in Norway's imports of goods. The countries are weighted according to their importance in Norwegian imports. Source: IMF (International Financial Statistics).
- $E$ = Import weighted exchange rate index (1991 = 1). The construction of the variable parallels that of $PF$. Source: Norges Bank (Norway's central bank).
- $U$ = Registered unemployment as a share of the labour force. Sources: Statistics Norway and the Directorate of Labour.
- $UC$ = Variable unit costs in the production of $X$. Source: Statistics Norway.

## References

Abbott, A. J. and Seddighi, H. R. (1996). 'Aggregate Imports and Expenditure Components in the UK: An Empirical Analysis', *Applied Economics* **28**, 1119-25.

Andersen, T. (1994). *Price Rigidity*, Clarendon Press, Oxford.

Anderton, B., Pesaran, B. and Wren-Lewis, S. (1992). 'Imports, Output and the Demand for Manufactures', *Oxford Economic Papers* **44**, 175-86.

Banerjee, A., Dolado, J. J., Hendry, D. F. and Smith, G. W. (1986). 'Exploring Equilibrium Relationships in Econometrics Through Static Models: Some Monte Carlo Evidence', *Oxford Bulletin of Economics and Statistics* **48**, 253-77.

Barker, T. (1977). 'International Trade and Economic Growth: An Alternative to the Neoclassical Approach', *Cambridge Journal of Economics* **1**, 153-72.

Barker, T. (1979). 'Identification of Activity Effects, Trends and Cycles in Import Demand', *Oxford Bulletin of Economics and Statistics* **41**, 63-68.

Beenstock, M. and Warburton, P. (1982). 'UK Imports and the International Trading Order', *Weltwirtschaftliches Archiv* **118**, 707-25.

Carlton, D. (1989). 'The Theory and Facts of How Markets Clear: Is Industrial Organization Valuable for Understanding Macroeconomics?', in Schmalensee, R. and Willig, R. D. (eds.), *Handbook of Industrial Organization,* Vol. 1, North-Holland, Amsterdam.

Carone, G. (1996). 'Modeling the U.S. Demand for Imports Through Cointegration and Error Correction', *Journal of Policy Modeling* **18**, 1-48.

Cheung, Y.-W. and Lai, K. S. (1993). 'Finite-Sample Sizes of Johansen's Likelihood Ratio Tests for Cointegration', *Oxford Bulletin of Economics and Statistics* **55**, 313-28.

Clarida, R. H. (1994). 'Cointegration, Aggregate Consumption, and the Demand for Imports: A Structural Econometric Investigation', *American Economic Review* **84**, 298-308.

Clarida, R. H. (1996). 'Consumption, Import Prices, and the Demand for Imported Consumer Durables: A Structural Econometric Investigation', *Review of Economics and Statistics* **78**, 369-74.

Cuthbertson, K. (1985). 'The Behaviour of UK Imports of Manufactured Goods', *National Institute Economic Review* **113**, August, 31-38.

Deaton, A. and Muellbauer, J. (1980). 'An Almost Ideal Demand System', *American Economic Review* **70**, 312-26.

de la Croix, D. and Urbain, J.-P. (1996). 'Intertemporal Substitution in Import Demand and Habit Formation', *Journal of Applied Econometrics*, forthcoming.

Doornik, J. A. and Hansen, H. (1994). A Practical Test for Univariate and Multivariate Normality, Discussion Paper, Nuffield College, University of Oxford.

Doornik, J. A. and Hendry, D. F. (1996). *GiveWin: An Interface to Empirical Modelling*, International Thomson Business Press, London.

Doornik, J. A. and Hendry, D. F. (1997). *Modelling Dynamic Systems Using PcFiml 9.0 for Windows*, International Thomson Business Press, London.

Drèze, J. H. and Bean, C. R. (eds.) (1990). *Europe's Unemployment Problem*, MIT Press, Cambridge, MA.

Engle, R. F. (1982). 'Autoregressive Conditional Heteroscedasticity with Estimates of the Variance of United Kingdom Inflation', *Econometrica* **50**, 987-1007.

Engle, R. F. and Granger, C. W. J. (1987). 'Co-Integration and Error Correction: Representation, Estimation and Testing', *Econometrica* **55**, 251-76.

Engle, R. F. and Hendry, D. F. (1993). 'Testing Superexogeneity and Invariance in Regression Models', *Journal of Econometrics* **56**, 119-39.

Engle, R. F., Hendry, D. F. and Richard, J.-P. (1983). 'Exogeneity', *Econometrica* **51**, 277-304.

Ericsson, N. R. and Irons, J. S. (1995). 'The Lucas Critique in Practice: Theory Without Measurement', in Hoover, K. D. (ed.), *Macroeconometrics: Developments, Tensions and Prospects*, Kluwer Academic Publishers, Boston, MA.

Favero, C. and Hendry, D. F. (1992). 'Testing the Lucas Critique: A Review', *Econometric Reviews* **11**, 265-306.

Gagnon, J. E. (1988). 'Adjustment Costs and International Trade Dynamics', *Journal of International Economics* **26**, 327-44.

Gerrard, W. J. and Godfrey, L. G. (1998). 'Diagnostic Checks for Single-Equation Error-Correction and Autoregressive Distributed Lag Models', *Manchester School* **66**, 222-37.

Goldstein, M. and Khan, M. S. (1976). 'Large Versus Small Price Changes and the Demand for Imports', *IMF Staff Papers* **23**, 200-25.

Goldstein, M. and Khan, M. S. (1985). 'Income and Price Effects in Foreign Trade', in Jones, R. W. and Kenen, P. B. (eds.), *Handbook of International Economics*, Vol. 2, North-Holland, Amsterdam.

Greenwald, B. and Stiglitz, J. E. (1989). 'Toward a Theory of Rigidities', *American Economic Review* **79**, 364-69.

Gregory, R. G. (1971). 'United States Imports and Internal Pressure of Demand: 1948-68', *American Economic Review* **61**, 28-47.

Haas, R. D. and Turner, G. (1990). 'The World Trade Model: Revised Estimates', *Journal of Policy Modeling* **12**, 93-128.

Hansen, B. E. (1992). 'Testing for Parameter Instability in Linear Models', *Journal of Policy Modeling* **14**, 517-33.

Hansen, H. and Johansen, S. (1993). Recursive Estimation in Cointegrated VAR-Models, Preprint 1, Institute of Mathematical Statistics, University of Copenhagen.

Harvey, A. C. (1981). *The Econometric Analysis of Time Series*, Philip Allan, Oxford.

Hendry, D. F. (1988). 'The Encompassing Implications of Feedback Versus Feedforward Mechanisms in Econometrics', *Oxford Economic Papers* **40**, 132-49.

Hendry, D. F. (1995a). 'On the Interactions of Unit Roots and Exogeneity', *Econometric Reviews* **14**, 383-419.

Hendry, D. F. (1995b). *Dynamic Econometrics*, Oxford University Press, Oxford.

Hendry, D. F. and Doornik, J. A. (1996). *Empirical Econometric Modelling Using PcGive 9.0 for Windows*, International Thomson Business Press, London.

Hendry, D. F. and Ericsson, N. R. (1991). 'Modeling the Demand for Narrow Money in the United Kingdom and the United States', *European Economic Review* **35**, 833-86.

Hendry, D. F. and Neale, A. J. (1988). 'Interpreting Long-Run Equilibrium Solutions in Conventional Macro Models: A Comment', *Economic Journal* **98**, 808-17.

Husted, S. and Kollintzas, T. (1987). 'Linear Rational Expectations Equilibrium Laws of Motion for Selected U.S. Raw Material Imports', *International Economic Review* **28**, 651-70.

Johansen, S. (1988). 'Statistical Analysis of Cointegration Vectors', *Journal of Economic Dynamics and Control* **12**, 231-54.

Johansen, S. (1991). 'Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models', *Econometrica* **59**, 1551-80.

Johansen, S. (1992). 'Cointegration in Partial Systems and the Efficiency of Single-Equation Analysis', *Journal of Econometrics* **52**, 389-402.

Johansen, S. (1995). *Likelihood-Based Inference in Cointegrated Vector Autoregressive Models*, Oxford University Press, Oxford.

Johansen, S. and Juselius, K. (1990). 'Maximum Likelihood Estimation and Inference on Cointegration – With Applications to the Demand for Money', *Oxford Bulletin of Economics and Statistics* **52**, 169-210.

Johansen, S. and Juselius, K. (1992). 'Testing Structural Hypotheses in a Multivariate Cointegration Analysis of the PPP and the UIP for UK', *Journal of Econometrics* **53**, 211-44.

Klein, L. R. and Summers, R. (1966). *The Wharton Index of Capacity Utilization*, Studies in Quantitative Economics No. 1, Department of Economics, University of Pennsylvania.

Kremers, J. J. M., Ericsson, N. R. and Dolado, J. J. (1992). 'The Power of Cointegration Tests', *Oxford Bulletin of Economics and Statistics* **54**, 325-48

Krugman, P. (1995). 'Growing World Trade: Causes and Consequences', *Brookings Papers on Economic Activity*, 1: 1995, 327-77 (with discussion).

Lächler, U. (1985). 'The Elasticity of Substitution between Imported and Domestically Produced Goods in Germany', *Weltwirtschaftliches Archiv* **121**, 74-96.

Liu, T. C. (1954). 'The Elasticity of U.S. Import Demand: A Theoretical and Empirical Reappraisal', *IMF Staff papers* **3**, 416-41.

Lucas, Jr., R. E. (1976). 'Econometric Policy Evaluation: A Critique', in Brunner, K. and Meltzer A. H. (eds.), *The Phillips Curve and Labor Markets*, North-Holland, Amsterdam.

Menon, J. (1995). 'Price and Activity Effects in International Trade: Cointegration, Aggregation and Prices', *Hitotsubashi Journal of Economics* **36**, 47-60.

Murray, T. and Ginmann, P. J. (1976). 'An Empirical Examination of the Traditional Aggregate Import Demand Model', *Review of Economics and Statistics* **58**, 75-80

Naug, B. and Nymoen, R. (1996). 'Pricing to Market in a Small Open Economy', *Scandinavian Journal of Economics* **98**, 329-50.

Neudorfer, P., Pichelmann, K. and Wagner, M. (1990). 'Unfavorable Shifts of Rationing Equilibria: Austrian Unemployment', in Drèze and Bean (1990), *op. cit.*

Orcutt, G. (1950).'Measurement of Price Elasticities in International Trade', *Review of Economics and Statistics* **32**, 117-32.

Osterwald-Lenum, M. (1992). 'A Note with Quantiles of the Asymptotic Distribution of the Maximum Likelihood Cointegration Rank Test Statistics', *Oxford Bulletin of Economics and Statistics* **54**, 461-72.

Perry, G. L. (1973) 'Capacity in Manufacturing', *Brookings Papers on Economic Activity*, 3: 1973, 701-42.

Rahbek, A. and Mosconi, R. (1998). The Role of Stationary Regressors in the Cointegration Test, Preprint No. 1, Department of Theoretical Statistics, University of Copenhagen.

Ramsey, J. P. (1969). 'Tests for Specification Errors in Classical Linear Least-Squares Regression Analysis', *Journal of the Royal Statistical Society* **31**, *Series B*, 350-71.

Reimers, H.-E. (1992). 'Comparisons of Tests for Multivariate Cointegration', *Statistical Papers* **33**, 335-59.

Sato, R. (1975). 'The Most General Class of CES Functions', *Econometrica* **43**, 999-1003.

Sato, R. (1977). 'Homothetic and Non-Homothetic CES Production Functions', *American Economic Review* **67**, 559-69.

Sedgley, N. and Smith, J. (1994). 'An Analysis of UK Imports Using Multivariate Cointegration', *Oxford Bulletin of Economics and Statistics* **56**, 135-50.

Temple, P. and Urga, G. (1997). 'The Competitiveness of UK Manufacturing: Evidence from Imports', *Oxford Economic Papers* **49**, 207-27.

Urbain, J.-P. (1992a). 'Error Correction Models for Aggregate Imports: The Case of Two Small and Open Economies', in Dagenais, M. G. and Muet, P.-A. (eds.), *International Trade Modelling*, Chapman & Hall, London.

Urbain, J.-P. (1992b). 'On Weak Exogeneity in Error Correction Models', *Oxford Bulletin of Economics and Statistics* **54**, 187-207.

Urbain, J.-P. (1995). 'Partial Versus Full System Modelling of Cointegrated Systems: An Empirical Illustration', *Journal of Econometrics* **69**, 177-210.

Urbain, J.-P. (1996). 'Japanese Import Behaviour and Cointegration: A Comment', *Journal of Policy Modeling* **18**, 583-601.

van Bergeijk, P. A. G. and Mensink, N. W. (1997). 'Measuring Globalization', *Journal of World Trade* **31**, 159-68.

Wallis, K. F. (1987). 'Time Series Analysis of Bounded Economic Variables', *Journal of Time Series Analysis* **8**, 115-23.

White, H. (1980). 'A Heteroskedasticity-Consistent Covariance Matrix Estimator and a Direct Test for Heteroskedasticity', *Econometrica* **48**, 817-38.

Wilson, J. F. and Takacs, W. E. (1979).'Differential Responses to Price and Exchange Rate Influences in the Foreign Trade of Selected Industrial Countries', *Review of Economics and Statistics* **61**, 267-79.

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