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HOW DOES THE SHARE OF IMPORTS CHANGE DURING STRUCTURAL ADJUSTMENT?

by

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ABSTRACT

Estimating the price responsiveness of market shares during a period of structural transition requires a distinction to be made between responses to variables explicitly recognized in the model and those due to more general changes in the trading environment. Often the latter are minimally modelled as market penetration curves taking the form of a sigmoid trend. Broadly this is the approach followed in the present paper; however, the trend 'parameter' capturing ultimate market share at a fixed level of price competitiveness is itself made a logistic function of the relative price variable measuring such competitiveness.

The application of the model is to quarterly data on the share of imports in Australian personal consumption over the 1980s and the first half of the 1990s. Most of the signal relevant to price competition between domestic and imported consumer goods occurred over the four years 1985–1988. This coincided with sizeable movements in the real exchange rate; and therefore, presumably, with collinear movements in the prices of the components within the domestic and the imported aggregates, which would be favourable circumstances for the application of Hicks' composite commodity idea. The responses in aggregate market shares during this episode suggest a very long-run Armington elasticity in the range 3.4 to 4.8, with short-run (quarterly) values of 0.6 to 0.8.

Keywords: import substitution, Armington elasticity, consumption, structural adjustment, logistic function, market penetration curve.

J.E.L. Classification numbers: D12, C13, C22, F47, 012.

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1. Introduction

Recent years have seen increasing market penetration by imports in several areas of the economy. If we regard this experience as reflecting mainly changes in Australian competitiveness – in other words, if we follow the Armington paradigm as implemented by Alaouze *et al.* [1] – we may come to the conclusion that substitution elasticities between domestic and foreign goods are very high, and therefore that even a slight deterioration in our competitiveness could lead to the annihilation of domestic import competing industries.

Is there an alternative view? This depends on how one interprets the current phase of global trading history. If the extensive relocation of production for manufactures in developing countries (especially in Asia) is seen as the movement from one equilibrium trading pattern to another, then the long-term outlook for Australia's import competitors may not be so grim. Whilst rapid displacement by imports may be occurring now, this could correspond to the rapid growth segment "A" of the market penetration curve shown in Figure 1.1.

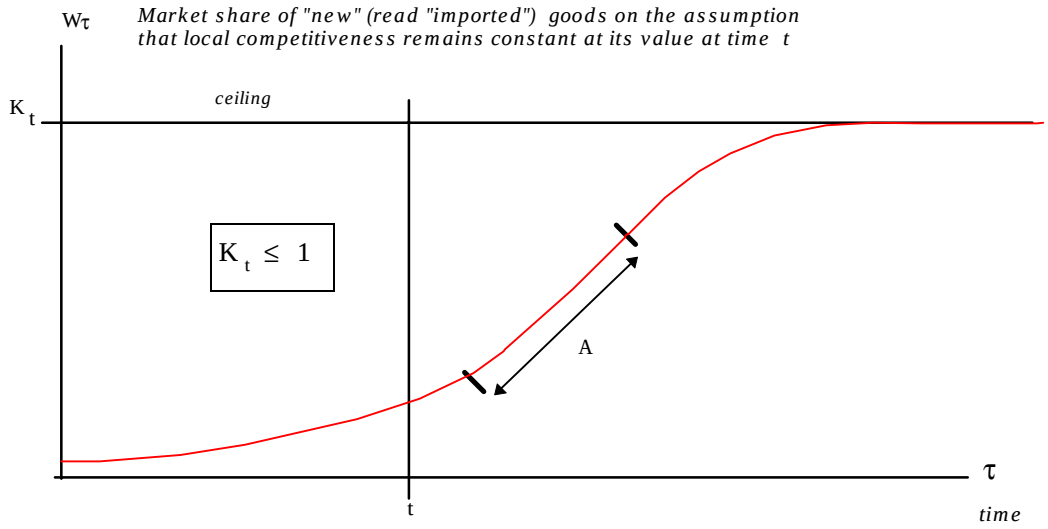

Figure 1.1 Market penetration curve. Note that a lowered competitiveness of local suppliers would lead to a higher ceiling on imports' final share of the market.

If the above view is accepted, then, as hazardous as all such exercises are, it leaves little alternative but to attempt to estimate the dynamics of the adjustment path.

* The author is grateful to Philip D. Adams for posing the problem, for supplying the initial and final data sets used in this paper, for helpful feedback and for his patience with me throughout; to Jayant Menon and Keith R. McLaren for useful discussions; and to Kaludura Abayasiri-Silva and Tim R. L. Fry for comments.

Some circumspection should accompany this exercise. As is obvious from Figure 1.1, if all sample data were to occur in the segment "A", and if there was relatively little time-series variation in competitiveness, then the location of the ceiling K_t would either be unidentified, or else determined largely by second-order parameters of the penetration curve. It is unlikely that we would be able, under the postulated conditions, to estimate such parameters with precision.

In fact in recent years there has been substantial variation both in competitiveness, and in market shares (at least for some major aggregates). In the balance of this paper the focus is on aggregate consumption in Australia. Imports' share (quantity index basis) and local competitiveness are shown in Figure 1.2. Given the variation evident in these data, it does seem worthwhile to attempt to disentangle the dynamics of market share adjustment from price responsiveness.

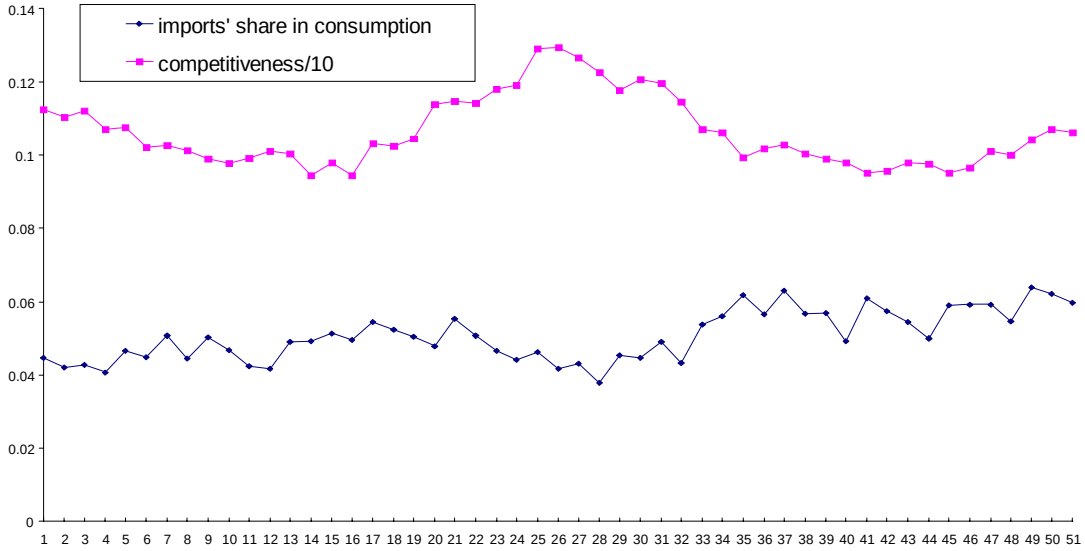

Figure 1.2 Share of imports in consumption and competitiveness index 1980Q3 through 1993Q1

The remainder of this paper is structured as follows. Section 2 contains a model with the following features:

(i) Imports' 'market share'[1] is a logistic function of time, with ultimate market penetration depending inversely on local competitiveness.

(ii) The ultimate market share of imports is itself a (displaced) logistic function of competitiveness.

Section 3 contains a discussion of the method used to estimate the model, while section 4 contains a brief discussion of the data. An initial set of results is given in Section 5. The implications for the performance of the model of a new data release are explored in Section 6, where a summary of the estimation results is also given. Finally, concluding remarks are offered in Section 7.

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1 'Market share' throughout this paper means the ratio of two quantities (rather than values). "Log" and "ln" are used interchangeably throughout this document to indicate natural logarithms.

2. The Double Logistic Model

2.1 The problem

The problem posed above may be restated succinctly as: given time-series data on imports' market share in consumption – about 13 years of quarterly observations – estimate the penetration curve subject to a sensible role for "competitiveness" (here defined as the ratio of the purchasers' prices of imported consumption goods to the prices of domestically produced consumption goods). Ultimately we want to forecast the (time) change in market share as a function of time and of change in competitiveness P_t, where

$$P_t \equiv P_{import}(t) / P_{domestic}(t) \quad .$$

For future reference, note $p_t \equiv d \ln P_t$.

2.2 Price responsiveness of the ceiling value of imports' share

In the basic model it is assumed that relative prices enter market dynamics through the 'ceiling', or saturation–share, parameter, K:

$$K_t = f(P_t) \; ; \quad 0 \leq K_t \leq 1 \quad . \tag{2.1}$$

Note (i) that K's dependence on time is solely via its dependence on competitiveness, P_t; and (ii) that P_t in (2.1) is playing the role of the long-run expectation held at time t for future competitiveness. More elaborate developments of the theory sketched below would replace P_t in (2.1) by such an expectational variable.

The form of f(•) in (2.1) is taken to be a displaced logistic in the logarithm of P_t:

$$K_t = \theta_{q(t)} + \frac{(1 - \theta_{q(t)})}{1 + \alpha_{q(t)} P_t^{\gamma}} \tag{2.2}$$

where q(t) is an indexing function which identifies the quarter of the year in which observation t falls, and the parameters are θ_1, θ_2, θ_3, θ_4 (the minimum ultimate share possible for imports), α_1, α_2, α_3, α_4 (location parameters) and γ (sensitivity of the ultimate share of imports to competitiveness). Other than γ, the parameters are indexed by quarter to account for seasonality in import patterns. The expected signs of the θs, αs and γ are positive. In addition, the θ–parameters are constrained to lie within the [0, 1] interval. A value of $\theta > 0$ implies that even if Australian competitiveness were to become superb ($P_t \to \infty$), imports would still ultimately achieve a share θ (albeit a small share) of the market.

In the derivations below it will be helpful to use a variable Q_t defined as

$$Q_t = \frac{1}{1 + \alpha_{q(t)} P_t^{\gamma}} \tag{2.3}$$

which is a pure logistic function with unit ceiling parameter. Then (2.2) may be rewritten as

$$K_t = \theta_{q(t)} + (1 - \theta_{q(t)}) Q_t \tag{2.4}$$

Notice that since the θs and Q_t are both bounded in [0, 1], then so is K_t. Also, since the θs are absolute constants, it follows that

$$\frac{\partial K_t}{\partial v} = (1 - \theta)\, \partial Q_t / \partial v \tag{2.5}$$

and that

$$\frac{\partial \ln K_t}{\partial v} = \frac{(K_t - \theta)}{K_t} \frac{\partial \ln Q_t}{\partial v} \tag{2.6}$$

where v is any parameter or variable, and where in omitting the q(t) subscripts on θ we have (for the time-being) neglected quarter-to-quarter changes due to the seasonal pattern in the θs. In particular,

$$\frac{\partial K_t}{\partial \gamma} = -(1 - \theta)\, \gamma\, \alpha_t\, Q_t^{\,2}\; P_t^{\,\gamma - 1} \;, \tag{2.7}$$

and

$$\frac{d \ln K_t}{dt} = -\gamma \frac{(K_t - \theta)}{K_t} (1 - Q_t) \frac{d \ln P_t}{dt} \;. \tag{2.8}$$

Notice that with $\gamma > 0$, (2.8) implies that the ultimate share of imports K_t is monotonic decreasing in Australian competitiveness P_t. This is illustrated in Figure 2.1.

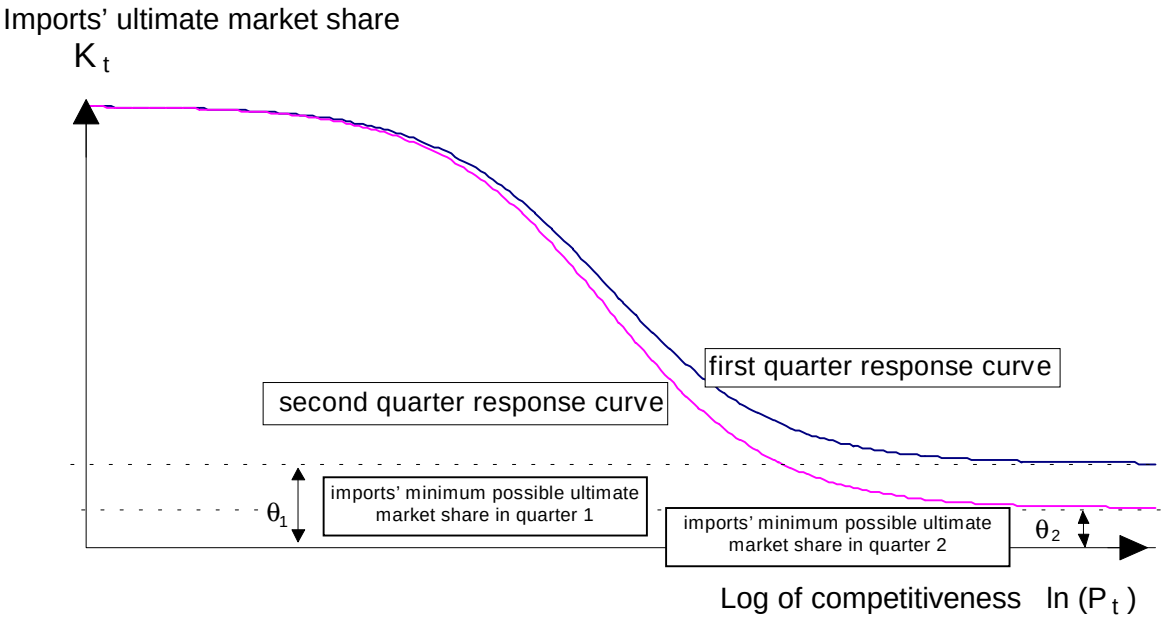

Figure 2.1 Decline in ultimate import penetration as Australia's competitiveness increases. Note that there is a different curve for each quarter, and possibly a finite positive lower bound on imports' ultimate market share which is quarter-specific.

2.3 Market adjustment through time — non-stochastic dynamics

With price-competition modelled via (2.2), a second logistic function is now introduced to model the sigmoid trend in market penetration shown in Figure 1.1:

$$W_t = \frac{K_t}{1 + a\, e^{bt}} \;, \tag{2.9}$$

where W_t is the share of imports (quantity basis) in total sales to consumption, and a and b are parameters with expected signs which are positive and negative respectively.[2]

The change over an interval of time dt in imports' market share can be found from:[3]

$$(2.10) \qquad d \ln \frac{(K_t - W_t)}{W_t} = b\, dt \ .$$

Using w_t and k_t to denote $d \ln K_t$ and $d \ln W_t$ respectively, (2.10) may be expressed as:

$$(2.11) \qquad w_t = k_t - \left(\frac{K_t - W_t}{K_t}\right) b\, dt \ .$$

If K is fixed, (2.11) implies that the logarithmic time rate of change of market share is directly proportional to the gap between the ceiling and the current share, where this gap is expressed as a fraction of the ceiling value, the constant of proportionality being (-b). For shares which grow with time it follows that ***b is negative***.

To simplify matters while ideas are clarified, we drop the seasonal variation in the θs and αs for the moment. Taking total differentials of (2.4) we obtain:

$$(2.12) \qquad k_t = \left(\frac{K_t - \theta}{K_t}\right) q_t \ ,$$

where $q_t \equiv d \ln Q_t$. Using (2.8) we obtain

$$(2.13) \qquad k_t = -\gamma \frac{(K_t - \theta)}{K_t} (1 - Q_t)\ p_t \ .$$

Substituting (2.13) into (2.11), we find

$$(2.14) \qquad w_t = \underbrace{-\gamma \left(\frac{K_t - \theta}{K_t}\right)(1 - Q_t)\ p_t}_{[term\ 1]} \underbrace{- b \left(\frac{K_t - W_t}{K_t}\right) dt}_{[term\ 2]} \ .$$

Consider a period in which competitiveness is rising ($p_t > 0$). Since $\alpha, \gamma > 0$, $b < 0$ and $K \geq W$, we see that the *proportional* change w_t in the share of imports at time t consists of two parts: a fall due to the effect of increased competitiveness *(term 1)* and a rise due to the working through of market penetration *(term 2)*. The first effect becomes positive if competitiveness falls; the second is always positive.

---

2 Note that we do not require W_t to approach θ as $t \to -\infty$; in fact, $W_t \to 0$ as $t \to -\infty$. This is interpreted to mean that zero is the natural origin for foreign market share before the structural shift in trading patterns.

3 To obtain this result, note that $(K - W_t)/W_t = a\, e^{bt}$; then take log differentials. Also note that the seasonality built into K via the specification of quarter-specific values of θ and α flows into W, implying that the ratio $(K_t - W_t)/W_t$ is free of seasonal influences and hence that the scalar dt on the right of (2.10) suffices – we are no longer neglecting the seasonal pattern in the θs.

Let M_t and Y_t denote imported and domestic 'quantities' (i.e., \$A of a base period). Then by definition we have

$$W_t \equiv M_t / (M_t + Y_t) \tag{2.15}$$

and

$$w_t \equiv (1 - W_t)(m_t - y_t) = (1 - W_t)\, d \ln \left(\frac{M_t}{Y_t} \right) \tag{2.16}$$

where m_t and y_t respectively are $d \ln M_t$ and $d \ln Y_t$.

Consider a fixed point of time and allow a change in competitiveness (in the sense of a deviation from control at t). From (2.14) and (2.16) we see that the elasticity of substitution between imports and the domestic good is[4]

$$\sigma_{SR} = - \left\{ d \ln \left(\frac{M_t}{Y_t} \right) \div d \ln \left(\frac{P_{import}}{P_{domestic}} \right) \right\} \Bigg|_{\textit{fixed utility level; t}} \tag{2.17}$$

$$= \gamma \left(\frac{K_t - \theta}{K_t} \right) (1 - Q_t) / (1 - W_t) = \gamma \left(\frac{K_t - \theta}{K_t} \right) \frac{(1 - K_t)}{(1 - \theta)} / (1 - W_t) \; .$$

Above σ carries subscripts SR to denote that it is a short-run value. In the long run, as $t \to \infty$, $W_t \to K_t$, and we have for the long-run substitution elasticity (at fixed P_t):

$$\sigma_{LR} = \gamma (K_t - \theta) / \{ K_t (1 - \theta) \} \; . \tag{2.18}$$

Since $W_t \leq K_t$, $\sigma_{LR} \geq \sigma_{SR}$. Notice that "long-run" in this case means that we compare the displacement between two situations in which market saturation conditional on P_t has occurred. Also notice that, as $\theta \to 0$, $\sigma_{LR} \to \gamma$, which justifies the description of γ as the *long-run sensitivity of imports' market share to local competitiveness*, more briefly referred to below as the price *sensitivity* parameter.

3. Estimation of the Model

The approach to estimation of the model is to list the desirable properties of the estimator, and to incorporate each, either within the objective function, or as a set of side constraints.

The properties considered desirable were the following:

(1a) The estimated model must never forecast a value of the share of imports that exceeds 1 or is less than zero.

(2a) If possible, the stochastic specification should be similarly constrained so that a realization of W_t outside the [0,1] interval is impossible.

(3a) The fitted $\{ W_t \}$ must provide within-sample unbiased estimates of the actual $\{ W_t \}$ in the sense that the sample regression of the actual share W_t on fitted values of this variable must be a straight line through the origin. (In the linear model, this is a property of OLS estimates.)

---

4 Notice that from (2.4) the ratio of M_t to Y_t at fixed t depends only on the price ratio P_t (via K_t). Thus the system behaves as through preferences are locally homothetic (i.e., at fixed t) and we do not need to constrain the differentials in (2.17) to be consistent with a fixed utility level since the latter is irrelevant for market share under our demand specification. Note, though, that imports' share can vary over time at a fixed value of competitiveness due to market penetration working itself through.

(4a) The fitted $\{\Delta W_t\}$ must be similarly unbiased, so that the sample regression of the actual changes in share ΔW_t on fitted values of this variable must also be a straight line through the origin.

(5a) The serial properties of the fitted equation must be satisfactory; in particular:

I. The right and left-hand sides of the fitted equation should pass at least informal tests of cointegration.

II. The first-order serial correlation of the residuals should be close to zero (with a Durbin-Watson statistic between 1.8 and 2.2, say). (Misspecification of seasonality would likely lead to violation of this requirement.)

III. Serial correlation between observations realized in the same season must be eliminated (with a 4th-order Durbin-Watson statistic between 1.8 and 2.2, say).

Subject to all of the above, the likelihood of the sample data should be maximized with respect to the parameters finally chosen.

The above requirements are incorporated into estimation as follows.

(1b) The logistic function (2.9) returns values of W_t which lie in $(0, K_t]$ for t in the interval $(-\infty, \infty)$, while the function (2.2) guarantees that K_t is bounded in (0, 1) provided that $\theta_{q(t)}$ is similarly bounded for each of its four possible values. We enforce the latter constraint in estimation. Using (2.9) as a forecasting equation will then satisfy (1a).

(2b) The requirement (2a) can be met by transforming the model using the log-ratio form (Fry, Fry and McLaren [1996].[5]

(3b) Requirements (3a), (4a) (5a)II and (5a)III are to be enforced directly in the estimation as side constraints on the optimization of the fit.

(4b) Requirement (5a)I is to be met by allowing for an error-correction mechanism in the fitted equation, and by visually checking the reversion to mean of the fitted equation.

(5b) Finally, the objective function is chosen to be the error sum of squares of the log-ratio transform of the model with error correction included. Under normality of the equation errors, the likelihood is maximized subject to the explicit constraints listed above.

3.1 Restricted maximum-likelihood estimator

It is assumed that the model is to be used to predict both the level and the quarterly changes in imports' market share, ΔW_1, ..., ΔW_{50}. Conditional on competitiveness P_t, the date t, and the parameters $\{\theta_1, \theta_2, \theta_3, \theta_4\ \alpha_1, \alpha_2, \alpha_3, \alpha_4, \gamma;\ a, b\}$, the projected share value at t, $\hat{W}_t$, is obtained from (2.9) (after substitution for K from (2.2)).

By construction, the $\hat{W}_t$ lie in the interval $[\theta_{q(t)}, K_t]$. It follows that the variable

$$\hat{V}_t = \frac{\hat{W}_t - \theta_{q(t)}}{K_t - \theta_{q(t)}} \tag{3.1}$$

satisfies the constraint

5 There may be some conflict with 5(a)I, however, because the log-ratio transform does not guarantee (2a) if an explicit error-correction term is needed in the finally fitted equation. As reported below, in the present study this turns out to be at worst a minor difficulty.

$$0 \leq \hat{V}_t \leq 1 . \tag{3.2}$$

Notice that if W_t indicates actual share data and V_t is defined by,

$$V_t = \frac{W_t - \theta_{q(t)}}{K_t - \theta_{q(t)}} \tag{3.3}$$

then, for arbitrary K_t it ***cannot*** be guaranteed that

$$0 \leq V_t \leq 1 ; \tag{3.4}$$

however, if K_t is restricted to be greater than the largest W_t observed in the sample, (3.4) will necessarily hold. It turns out that our estimates below respect the inequality

$$\hat{K}_t \geq \underset{t \in \textit{sample}}{\text{Max}} \{ W_t \} . \tag{3.5}$$

Application of the log-ratio transform (Fry, Fry and McLaren [1996]) to this model involves forming the variables

$$\hat{L}_t = \hat{V}_t / (1 - \hat{V}_t) \tag{3.6a}$$

and

$$L_t = V_t / (1 - V_t) . \tag{3.6b}$$

The basic idea underlying this transform is to map the variables V_t and $\hat{V}_t$, which both are defined on the closed unit interval [0, 1] onto the real line $(-\infty,\infty)$. The model could be fitted in the levels as

$$\ln L_t = \ln \hat{L}_t + v_t , \tag{3.7}$$

or in the differences as

$$\Delta \ln L_t = \Delta \ln \hat{L}_t + u_t , \qquad [\Delta(\bullet)_t \equiv (\bullet)_{t+1} - (\bullet)_t] \tag{3.8}$$

where v_t and u_t are zero-mean random errors (with $u_t \equiv \Delta v_t$). Under the conditions postulated above, the values of W_t estimated from (3.7) are guaranteed to remain in the [0, 1] interval, while the domain of variation of u_t remains unrestricted.

Given the widespread problem of high serial correlation, it used to be common to choose the difference form of the equation for estimation; the critique of the cointegrationists that such a procedure throws away the long-run relationship of interest latterly has caused empirical workers to think twice about differencing.

So far as the current paper is concerned, this part of the critique clearly is not applicable, since the parameters of the long-run relationships imbedded in {(2.3), (2.4), (2.9)} are recoverable from the estimation of either (3.7) or (3.8). The argument against first differencing must then proceed in terms of efficiency of estimation, rather than lack of identification.

3.2 Error correction

In models with loosely specified deterministic parts, error correction mechanisms (ECMs) are usually needed in difference equations because no other device is present to ensure that the model converges over time to a stable and interpretable long run. As we have seen above, in the present case this does not apply. Nevertheless, there is a good reason for adding an ECM to the right of (3.8) — it can provide a benchmark to

assess whether the process as modelled in (3.8) is successful is ensuring that the estimated process reverts to mean. If this is the case, the ECM should have a coefficient close to zero (which can be checked from the estimates). For this reason we replace (3.8) for estimation purposes by:

$$\Delta \ln L_t = \Delta \ln \hat{L}_t + \psi \ln \left(\hat{L}_t / L_t \right) + u_t \ . \qquad (\psi > 0) \tag{3.9}$$

Alternatively, we can think of (3.9) as the Cochrane-Orcutt transform of (3.7) with autoregressive parameter $\rho \equiv 1 - \psi$. Even if $\psi = 0$, $\hat{V}_t$ and V_t may remain cointegrated in the sense that the internal logistical structure imbedded in both may stop them drifting apart over time.

A more compact notation for (3.7) and (3.9) is:

$$Y_t = X_t + v_t \ , \tag{3.10}$$

and

$$y_t = x_t + \psi \left(X_{t-1} - Y_{t-1} \right) + u_t \ . \tag{3.11}$$

in which the variables x and X are functions of the data and the parameters of the model (with $X = \ln \hat{L}$, $Y = \ln L$, $x = \Delta \ln X$ and $y = \Delta \ln Y$). Note that the dependency of y and Y on data is via realizations of the endogenous variable W_t only, whereas that of x and X is via exogenous variables only.

3.3 Stochastic specification

We assume throughout that the errors v and u in (3.7) and (3.8) have zero means. If v follows a first-order Markov scheme (i.e., is AR(1)) with a first-order serial correlation equal to (1 - ψ), then (3.9) is the appropriate equation for estimation and should yield a Durbin-Watson statistic close to 2. On the other hand, if v is close to being serially uncorrelated, then (3.7) is the appropriate estimating equation and should yield residuals with a DW close to 2; the errors u in (3.8), on the other hand, will follow a first-order moving average (MA(1)).

The small sample size (N=51) and the relatively large number of parameters (thirteen including the error variance) make it unlikely that powerful discrimination between

H_A: $v_t \sim AR(1)$ & u_t classically well behaved	and	H_B: $u_t \sim MA(1)$ & v_t classically well behaved

would be possible were the model not so highly constrained. Below attempts are made to estimate the model under H_A and under H_B. Unlike the more conventional approach to estimation, the maintained hypotheses are to be 'enforced' rather than tested. Thus under H_A the objective function is to be optimized in that part of the parameter space yielding diagnostics indicating freedom from serial correlation of the fitted u_ts; under H_B, optimization is to take place in the region that yields diagnostics indicating freedom from serial correlation in the fitted v_ts.

3.4 Objective function

Under H_A, the objective function is the sum of squares,

$$(3.12) \qquad \phi_A = \sum_{t=1}^{51} \hat{u}_t^2 ;$$

under H_B, the objective function is

$$(3.13) \qquad \phi_B = \sum_{t=1}^{50} \hat{v}_t^2 .$$

In the last two equations, the hats (^) indicate the fitted values of the equation errors. Thus our estimator is based on (restricted) non-linear least-squares applied to (3.11) in the case of H_A, and to (3.10) in the case of H_B. Under normality of the error distributions, these estimators (with minimands ϕ_A and ϕ_B) are restricted maximum-likelihood estimators under their respective maintained hypotheses H_A and H_B.

3.5 Side constraints

Unrestricted maximum likelihood will find parameters that yield values of $\Delta\hat{W}_t$ that are biased estimates of ΔW_t even within the sample. This is avoided by imposing two side constraints on the minimization of (3.12); namely,

$$(3.14a) \qquad \sum_{t=1}^{50} \Delta\hat{W}_t^* \, \Delta W_t^* \; / \sum_{t=1}^{50} \Delta[\hat{W}_t^*]^2 = 1 ,$$

and

$$(3.14b) \qquad \text{SAMPLE MEAN}\{\Delta\hat{W}_t\} = \text{SAMPLE MEAN}\{\Delta W_t\}$$

where the asterisks in (3.14a) indicate that the difference series have been expressed as deviations from their respective sample means. Similarly, to correct bias in the levels, a second pair of side constraints are imposed:

$$(3.15a) \qquad \sum_{t=1}^{51} \hat{W}_t^* \, \hat{W}_t \; / \sum_{t=1}^{51} [\hat{W}_t^*]^2 = 1 ,$$

and

$$(3.15b) \qquad \text{SAMPLE MEAN}\{\hat{W}_t\} = \text{SAMPLE MEAN}\{W_t\}$$

The left of (3.14a) [or (3.15a)] will be recognized as the slope of the OLS regression of ΔW_t [or W_t] on $\Delta\hat{W}_t$ [or $\hat{W}_t$]; a slope of unity will imply $\Delta\hat{W}_t$ [or $\hat{W}_t$] is an unbiased estimate of ΔW_t [or W_t] provided the intercepts in the aforementioned regressions are zero; the latter is guaranteed by (3.14b) and (3.15b).[6]

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6 'Keep in mind that here 'unbiased' is a sample concept: it simply means (as in the case of the ordinary-east-squares estimator of a linear equation) that the slope of the fitted regression of the actual values of the regressor on the fitted values is one; and that this fitted regression goes through the origin.

4. The Data

The data used in the estimation were taken from quarterly national accounts and balance of payments data obtained from the Australian Bureau of Statistics (ABS) on floppy disk. These data are consistent with the March quarter 1993 releases of the ABS publications with catalogue numbers 1.0.5206.0 and 5302.0. All data were seasonally unadjusted.

The sample period used for estimation consists of the 51 quarters 1980Q3 through 1993Q1.

The import-share data for W_t were derived from the series for total real private consumption (from the national accounts) and the series for real imports of consumption goods (from the balance of payments). Both series were expressed in terms of values at constant 1989-90 prices.

Data for the relative-price variable P_t were calculated from series for the implicit price deflator for private consumption (from the national accounts) and the implicit price deflator for imports of consumption goods (from the balance of payments). A series for the price deflator for domestically produced goods consumed by households was derived by assuming, for quarter t,

$$(4.1) \qquad P_{domestic}(t) \quad = \quad \left(\frac{P_{ipc}(t) \; - G_{impv}(t)\, P_{import}(t)}{G_{domv}(t)} \right),$$

where $G_{impv}(t)$ and $G_{domv}(t)$ are current-value weights of imports and locally produced goods in consumption, while $P_{ipc}(t)$ and $P_{import}(t)$ are the price deflators available directly from the ABS data on total consumption and on imports.

Data for $G_{impv}(t)$ and $G_{domv}(t)$ were deduced from ABS national-accounts data on nominal total household consumption expenditure and on balance-of-payments data on the value at current prices of imports of consumption goods.

Newly released data and their implications for this study are discussed in Section 6.

5. Results

The problem was set up in Excel 5.0 on a Pentium personal computer. The non-linear optimization was carried out using that package's powerful *Solver* routine. The constraints (3.14a&b), (3.15a&b) were imposed initially. It turned out that some relaxation of some of them (detailed below) was necessary to obtain feasible solutions.

5.1 Estimates under the autoregressive hypothesis H_A

As explained above, initially the estimation was set up to search only over the subset of the parameter space that produces acceptable diagnostics on the serial properties of the residuals. It turned out that the unbiasedness constraints and the constraints on the DW statistics could not simultaneously be met by the data. The lower bounds for the DW constraints **(5a) II&III** were relaxed from 1.8 to 1.7; when it still proved impossible to obtain a feasible solution, these bounds were further relaxed and eventually abandoned, resulting in a fourth-order DW as low as 1.53.[7] The resulting point estimates are shown in Table 5.1. Only one constraint was binding at solution: its purpose was to ensure that the slopes involved in **(3a)** and **(4a**) were both close to unity. Details are shown in Table 5.2.

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7 In the non-linear world of this model, no guidance is available on the bounds of the significance points of the 4th-order DW. Extrapolating in an informal manner from Giles and King's (1978) tabulations suggests that with 51 observations, a DW of about 1.5 would probably fall in the indeterminate range.

Table 5.1

Restricted Maximum Likelihood Estimates of the Double Logistic Model fitted in the First Differences under Hypothesis H_A
Quarterly Data, 1980Q3–1993Q1

Parameters of displaced logistic for the saturation share of imports as a function of Australian competitiveness

	1st quarter	*2nd quarter*	*3rd quarter*	*4th quarter*	*Price Sensitivity* γ
Location parameter	α_1	α_2	α_3	α_4	4.823
	0.1069	0.2043	0.1186	0.1202	
Minimum market share for imports:	θ_1	θ_2	θ_3	θ_4	
	0.0175	0.0003	0	0.0100	

Parameters of logistic trend in imports' market share		Error correction parameter in eqn (3.11)	Long- and short-run Armington elasticities (a)	Value of objective function (b)	Durbin-Watson (c)	Quasi-R^2 (levels) (d)	Quasi-R^2 (differences) (e)	Quasi-R^2 (differences) (f)
a	b	ψ	σ_{LR} at sample end 4.811	ϕ_A = 0.3597	DW(1) 2.296	QR^2(W)	QR^2(ΔW)	QR^2(Δ ln L)
18.039	-0.00525	0	typical σ_{SR} 0.790	[ϕ_B = 1.01×10^{-3}]	DW(4) 1.526	0.758	0.559	0.933

(a) σ_{LR} is calculated using (2.18); σ_{SR} is the sample mean of values calculated using (2.17).

(b) ϕ_A is the residual sum of squares from the finally fitted equation (3.9).

(c) DW(1) and DW(4) are the 1st and 4th-order Durbin-Watson statistics; they refer to residuals from the fitted equation (3.9). Corresponding statistics for (3.10) are 0.972 and 1.554.

(d) QR^2(W) is the square of the simple correlation coefficient between $\hat{W}_t$ and W_t.

(e) QR^2(ΔW) is the square of the simple correlation coefficient between $\Delta\hat{W}_t$ and ΔW_t.

(f) QR^2(Δ ln L) is the square of the simple correlation coefficient between $\Delta \ln \hat{L}_t$ and $\Delta \ln L_t$ from fitted equation (3.19).

Formula (2.18) indicates that when θ is small relative to K, the long-run substitution elasticity σ_{LR} is well approximated by the price sensitivity γ. Hence there is very little variation over time in the long-run substitution elasticity for the two quarters (namely, the second and third) in which imports' notional minimum share (at superlative levels of Australian competitiveness) is effectively zero in Table 5.1. The error correction parameter ψ is estimated to be zero to machine accuracy, thus favouring H_A and the first-difference form (3.8) as the appropriate vehicle for estimation.

Table 5.2

Constraint Binding on Parameter Solution Presented in Table 5.1

	Reference/ original form of constraint	Final value of coefficient	*Final form of constraint/* [value of Lagrange multiplier]
3(a)	slope of sample regression of W on $\hat{W} = 1$	1.00000	*average over the 2 regressions of the absolute deviation from 1 of the regression slope* $\leq 10^{-5}$
4(a)	slope of sample regression of ΔW on $\Delta\hat{W} = 1$	1.00002	[-1.1103]

The error-correcting tendency of the model (which, with $\psi = 0$, lacks an explicit error-correction term) can be assessed (at least informally) from Figure 5.1, which shows the 'actual' values of L (i.e., the values computed from the realized values of the endogenous shares and from the fitted parameters) and the fitted values $\hat{L}_t$ (computed from the exogenous variables and the fitted parameters). At least within sample, there seems to be little tendency for the fitted values to diverge from the 'observed' series as time progresses.

The seasonality of the data becomes very prominently displayed in the transform L. Figure 5.1 suggests that the parameterization on quarter-specific minimum shares θ and location parameters α successfully accounts for the seasonality. The estimates imply that only in the first and fourth quarters is imports' minimum possible ultimate share of consumption detectably above zero ($\theta_{1(t)} = 0.017$; $\theta_{4(t)} = 0.010$; or 1.7 and 1 per cent of the market respectively).

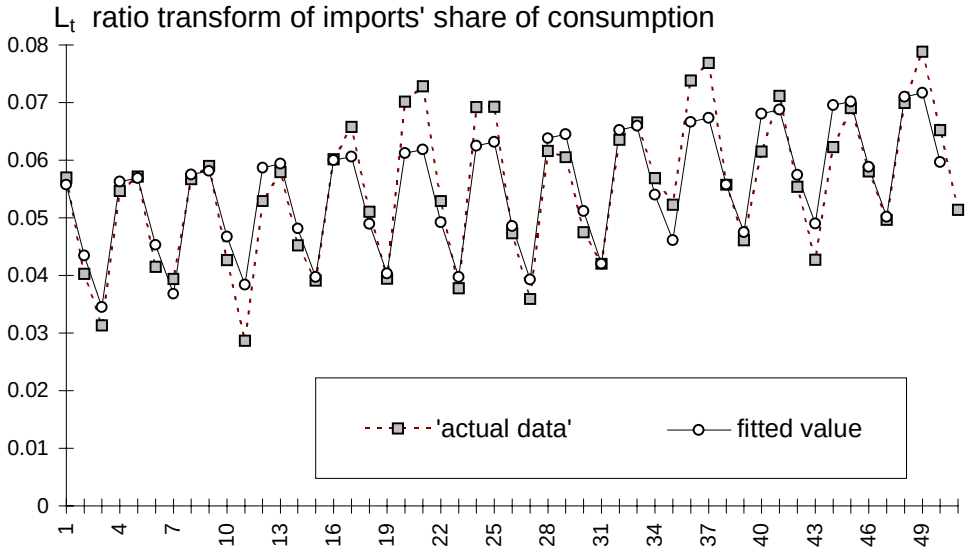

Figure 5.1 Actual and fitted values of the ratio transform of imports' market share, 1980Q3–1993Q1

Of more interest to the economist is the plot of the fitted and actual values of imports' market share shown in Figure 5.2

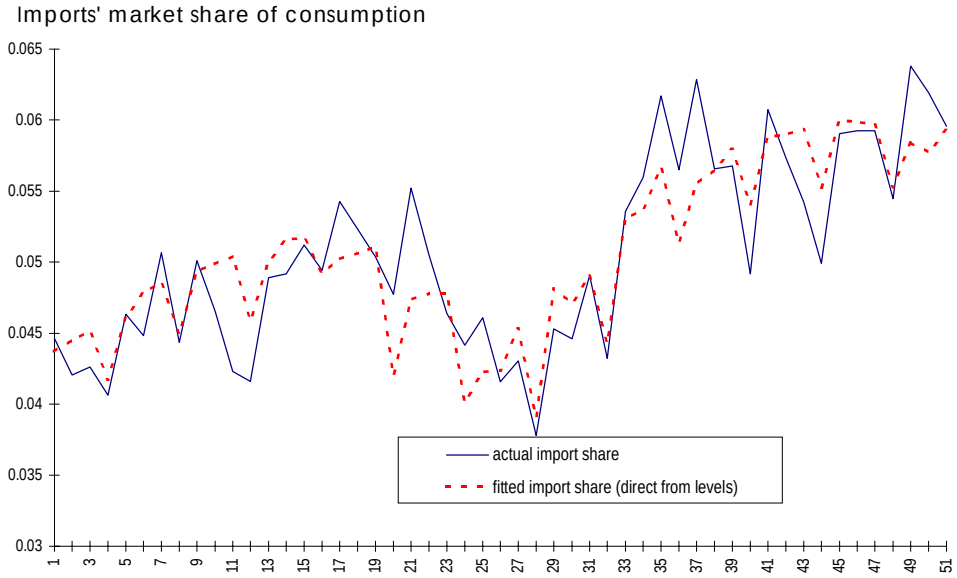

Figure 5.2 Projected and actual shares of imports in consumption, 1980Q3—1993Q1(estimated under H_A)

Also of interest is the ability of the model to project changes in imports' market share. The fitted and actual values of ΔW are shown in Figure 5.3.

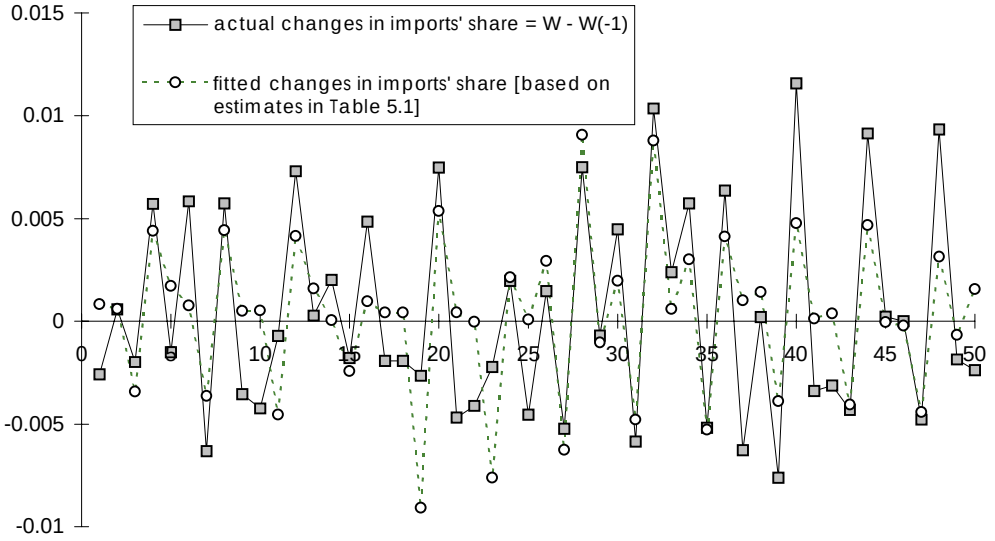

Figure 5.3 Projected and actual changes in shares of imports in consumption, 1980Q3—1993Q1(estimated under H_A)

The actual and fitted values of the log-ratio transform $\Delta \log L$ are charted in Figure 5.4, and the residuals from the fitted equation (3.9) are shown in Figure 5.5.

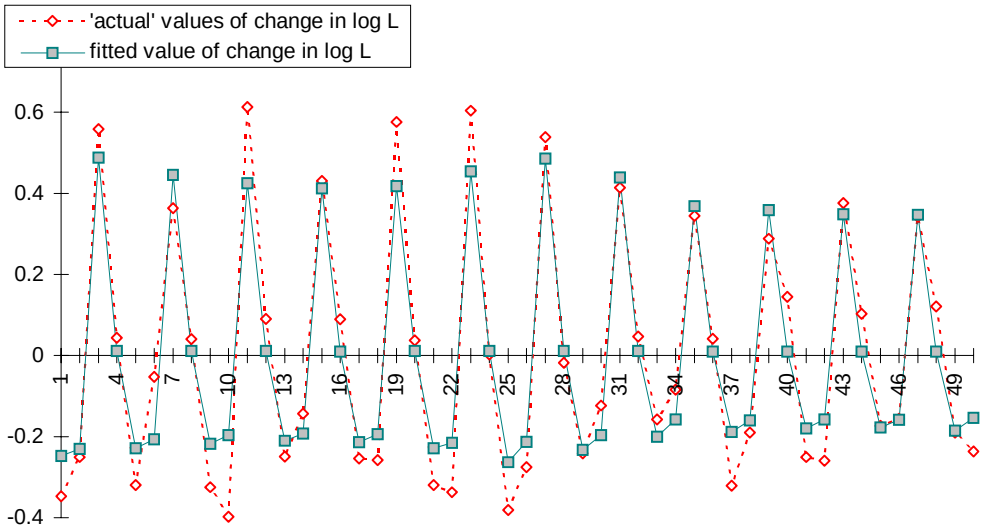

Figure 5.4 Projected and 'actual' values of Δ log L from equation (3.9), 1980Q3–1993Q1

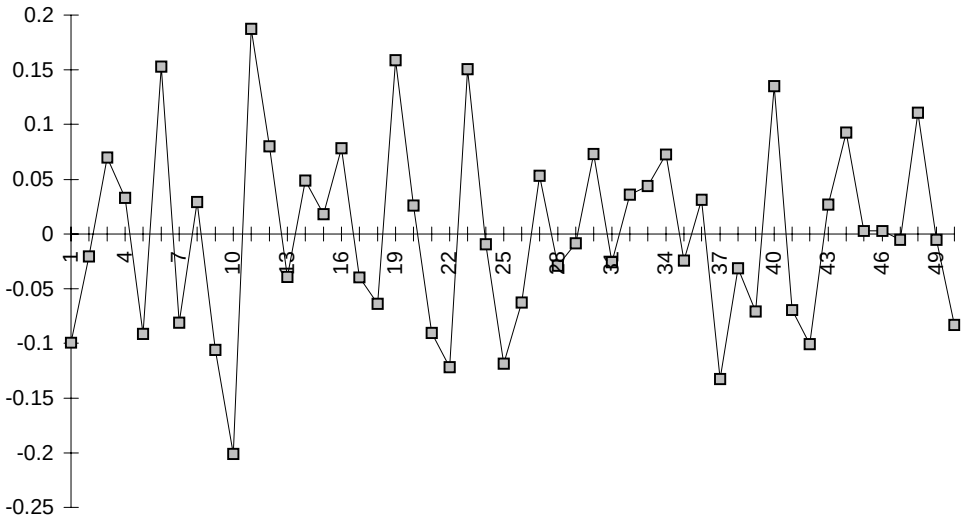

Figure 5.5 Residuals from the fitted equation (3.9), 1980Q3–1993Q1

How confident can we be that the price responsiveness captured by our estimated γ is genuine? Formal tests of significance cannot be used for several reasons, including:

(I) The sampling properties of the estimator used here are unknown.

(2) The sample size (51) is very small relative to the number of parameters (12) estimated.

(3) The specification of the model was developed interactively with data exploration.

Nevertheless, we can be reasonably sure that γ differs significantly from zero. At least this is so if one accepts the basis on which the estimator was constructed (unbiasedness within sample, and relative freedom from serial correlation in the residuals). This is established by searching over the parameter space for the set

which minimizes the squared value of γ subject to the constraint set imposed in the earlier estimation.

In principle such a search could return the value 0 for γ. It seems that the nearest to 0 that γ can go without leaving the feasible set is 2.96 (versus 4.82 in Table 5.1).[8] The lower role for substitutability in this parameter set results in a somewhat larger role for pure trend (b = -0.0065 versus -0.0053 in Table 5.1). Ignoring the constraint set and putting γ to zero, but keeping all other parameter values as in Table 5.1, causes the quasi-R^2 for the level of imports' market share to drop from 0.758 to 0.590; the fit for the differences actually improves from O.559 to 0.607.

When $\gamma = 0$, the model reduces to a (displaced) logistic time trend with quarter-specific upper and lower asymptotes. If constraints other than those for unbiasedness and those requiring the αs and θs to be non-negative are dropped, and ϕ_A is minimized subject to $\gamma = 0$, the model completely fails to track the dip below trend in imports' market share between quarters 22 and 32. This is shown in Figure 5.6. As shown in Figure 5.2, the results in Table 5.1 pick up this dip because of the coincident improvement in competitiveness (see Figures 1.2 and 6.2).

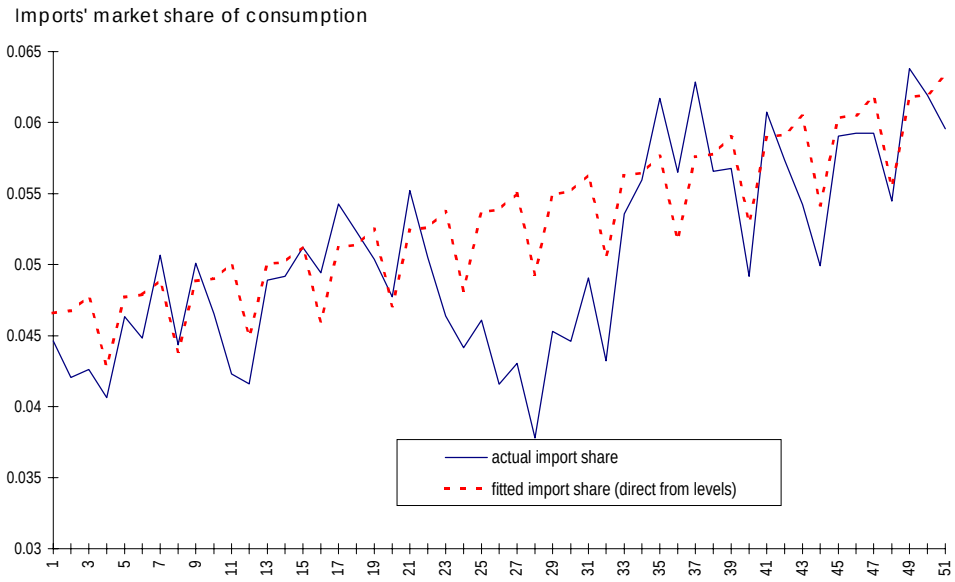

Figure 5.6 Fit of model, 1980Q3–1993Q1, when price influences are removed (i.e., $\gamma = 0$)

8 Here 'feasible set' excludes the DW restrictions, but includes the unbiasedness restrictions. The estimates of the parameters and fit statistics obtained are as follows: $\hat{\alpha}$ = (0.1058, 0.2206, 0.1151; 0.1176); $\hat{\gamma}$ = 2.959; $\hat{\theta}$ = (0.0176, 0.0003, 5.622×10^{-5}, 0.0010); a = 18.653; b = –0.00650; σ_{LR} (sample end) = 2.952; typical σ_{SR} = 0.435; ϕ_A = 0.3631; ϕ_B = 0.00109; DW(1)[differences] = 2.182; DW(4)[differences] = 1.538; QR^2(W) = 0.750; QR^2(ΔW) = 0.619; QR^2(Δ ln L) = 0.939; DW(1)[levels] = 0.728; DW(4)[levels] = 1.617.

For the reasons listed above, we cannot put a confidence band around our estimate of b, the trend coefficient. We note, however, that the estimate of this parameter is relatively stable under the variations in estimation approach explored in this paper, consistently yielding values in the range -0.005 to -0.007. With competitiveness completely eliminated from the story, as in Figure 5.6, the estimated value of b is -0.00625.

Can we say anything about the cointegration (or lack thereof) of equation (3.9)? In a formal sense, no. However, without the need for error correction, there seems to be no tendency within sample for $\Delta \ln L$ and $\Delta \ln \hat{L}$ to drift apart over time (Figure 5.4) ; the same is true of L and $\hat{L}$ (Figure 5.1). Perhaps more importantly, in the space of the primary data, W and $\hat{W}$ do not seem to be mutually diverging (Figure 5.2), nor do ΔW and $\Delta \hat{W}$ (Figure 5.3).

5.2 Estimates under the moving-average hypothesis H_B?

It did not prove possible to satisfy simultaneously the unbiasedness criteria **(3a)** and **(4a)** and obtain acceptable serial properties for the residuals from the levels equation (3.7). Results obtained with the unbiasedness restrictions, but not the DW restrictions, are virtually identical[9] to those reported in Table 5.1, and are not reported here.

5.3 Interpretation of results, I — the bad news

The bad news suggested by these results is that the ultimate penetration of the consumer goods market by imports is potentially very high: the values K_t of imports' ultimate market share at levels of competitiveness observed in the sample never fall below 71 percent in any of the estimates reported above. The time-series of the estimated within-sample values of K is shown in Figure 5.7.

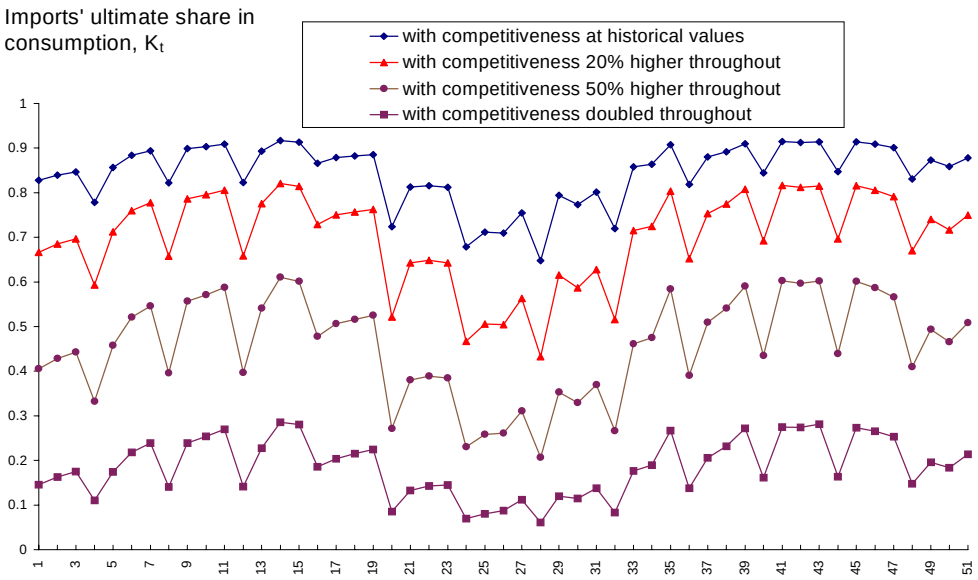

Figure 5.7 Estimated values of imports' ultimate share of consumption at levels of competitiveness that obtained during the sample, 1980Q3–1993Q1 (old data)

9 "Virtually identical" means identical to three or more significant digits. The plots of the fitted curves (as in Figures 5.1-5.5) cannot be resolved.

This news remains gloomy if one asks the question: by how much would competitiveness have had to increase in order to have lowered K by 30–40 percentage points throughout the sample period? This question roughly corresponds to the second lowest time path in Figure 5.7, which assumes that competitiveness throughout the sample were 50 percent higher than the historical values.

5.4 Interpretation of results, II — the good news

The above picture predicts a very substantial rise in imports' *ultimate* share of the market *even if Australian competitiveness were to reattain, and maintain, its highest sample value.* The good news is that the trend coefficient b is relatively small. Using formula (2.11) we can deduce the approximate length of time which must elapse at a given level of competitiveness for the local product to lose (say) ten percentage points of market share:

$$[dt]_{10} = 0.1\left\{\left(\frac{K_t - W_t}{K_t}\right) b\, W_t\right\}^{-1} . \tag{5.2}$$

At the end of the sample, this period is about 80 years if we work from Table 5.1, and slightly less if we use Table 5.3. Correcting for the linearization error inherent in (5.2) brings these values down to about 50 years. The displacement of the local product may be inexorable, but it is not rapid!

5.5 Interpretation of results, III — short-run price substitutability

Short-run elasticities of substitution vary over time at a fixed level of competitiveness because W_t on the far right-hand side of (2.17) does so. The estimated within-sample substitution elasticities $\sigma_{SR}(t)$ are shown in Figure 5.7. The plot reveals wide variation from about 0.3 through about 1.8, with the highest values occurring at the historically highest levels of competitiveness.[10]

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[10] The reason for the positive relationship between $\sigma_{SR}(t)$ and competitiveness P_t can be found by taking the total logarithmic differential of (2.17) and substituting for w_t from (2.11). We find that, at fixed t, $\sigma_{SR}(t)$ rises with K_t if and only if $\left\{ \frac{\theta_t - K_t^2}{(K_t - \theta_t)(1 - K_t)} + \frac{W_t}{1 - W_t} \right\} > 0$. The sample values of $\left\{ \frac{\theta_t - K_t^2}{(K_t - \theta_t)(1 - K_t)} + \frac{W_t}{1 - W_t} \right\}$ (where the Ws are the fitted values of imports' share) lie in the interval [-15.9, -2.3]. Hence throughout the sample, $\sigma_{SR}(t)$ *falls* with rises in K_t. Given the negative relationship between K_t and P_t implied by (2.2), it follows that $\sigma_{SR}(t)$ at each t rises with competitiveness P_t.

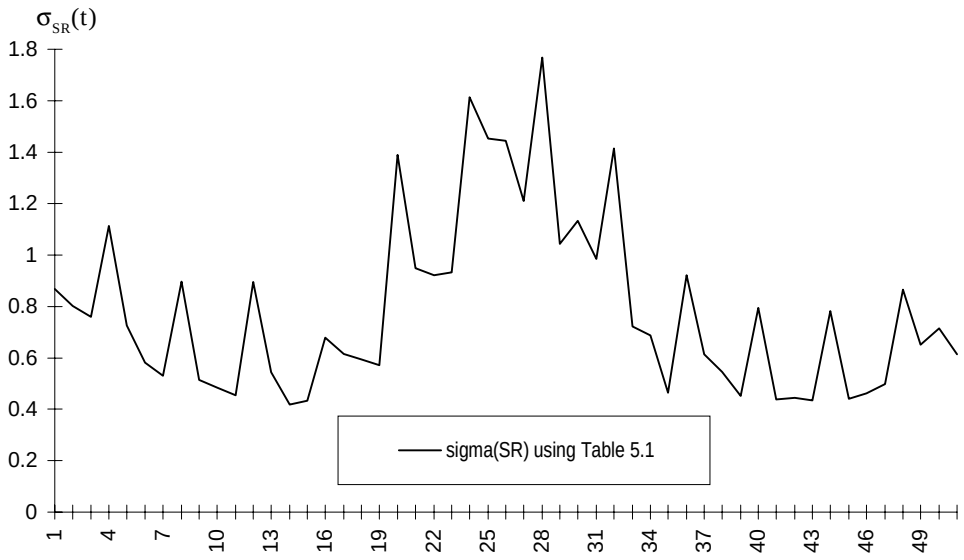

Figure 5.8 Estimated values of short-run substitution elasticity between imported and domestically produced consumption goods, 1980Q3 through 1993Q1

6. New data release, revised estimates and post-sample tracking

The results reported in Section 5 used the data available in the middle of 1993 (when this study, conducted intermittently among other work, began). The above work was completed during the first quarter of 1996. By then new data were available. This presented an opportunity for post-sample evaluation of the model. Unfortunately, though, the more recent ABS data (national accounts and balance of payments data including the December quarter of 1995) have been rebased and come on a new set of definitions. Thus the post-sample data and the within-sample data are not readily compared.

Data are available on both bases, however, from 1981Q3 through 1993Q1. This spans all but the initial four quarters of the sample which produced the results in Section 5. The data on imports' share in consumption and on competitiveness on both bases is plotted in Figures 6.1 and 6.2 respectively.

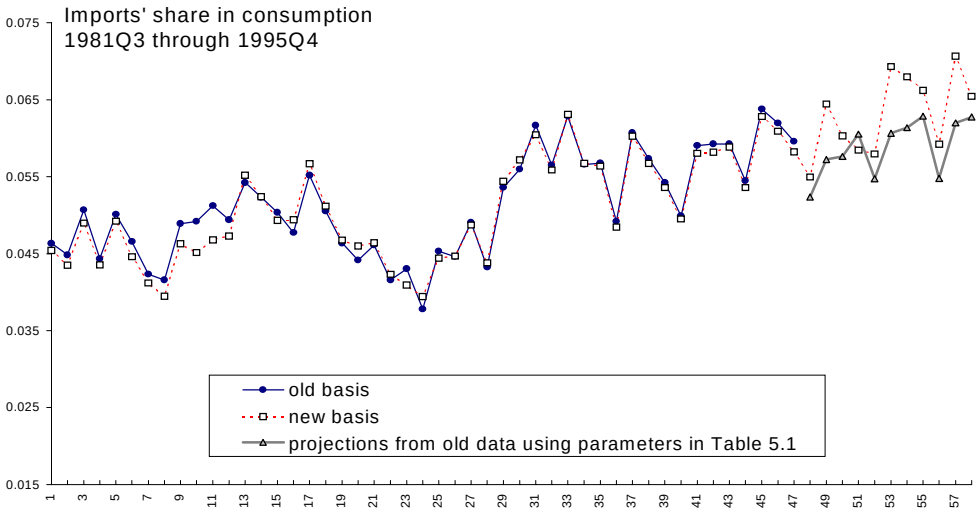

Figure 6.1 Data on imports' share in consumption on old and new basis, with projections from old estimates on new basis, 1993Q2–1995Q4

Figure 6.1 also shows the projections for the eleven quarters 1993Q2–1995Q4 obtained using the estimates in Table 5.1. Other than time, competitiveness is the exogenous variable driving the share projections. The latter are too low in ten out of the eleven quarters; part of the problem seems to be the different pattern of seasonality in the new data relative to the old — the seasonal peak in θ occurs in the September quarter in the new data, but in the March quarter in the old.

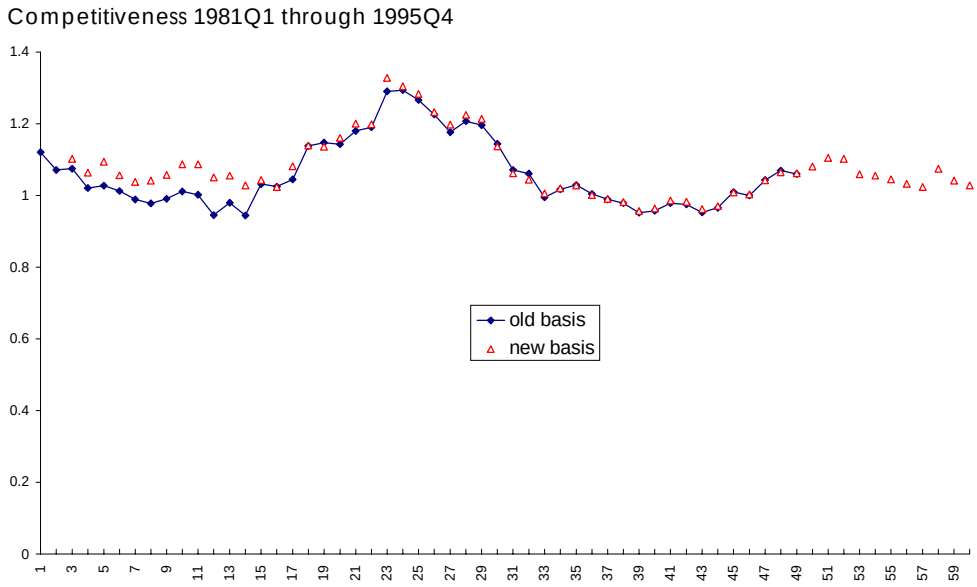

Figure 6.2 Data on competitiveness on old and new basis

Using the previously estimated parameter values as starting point, the model was re-estimated using the new data set. As before, it turned out that the restrictions on the serial properties of the residuals and on unbiasedness could not be met simultaneously by these data. The restrictions on the Durbin-Watson statistics were dropped, yielding sample values of 2.37 and 1.33 at first and fourth order respectively.

One remaining constraint was binding at solution: namely, the one to keep the slopes of the regressions of W and ΔW on their respective fitted values close to unity (as in Table 5.2). The resultant estimates are shown in Table 6.1. Unfortunately, these are not the restricted MLEs; as we will see below, $\gamma = 0$ yields a higher value of the likelihood function. However, for reasons which will become clear presently, the estimates corresponding to the relative maximum reported in Table 6.1 are not without interest.

Fits of W in the levels and the differences are shown in Figures 6.3 and 6.4, while the fits of the differences of the log-ratio transform and the associated residuals are shown in Figures 6.5 and 6.6 respectively. All of these charts are based on Table 6.1.

With γ set to zero, the model becomes a simple logistic trend with seasonal adjustment. The estimated parameters of this pure trend model are shown in

Table 6.1

*Estimates of the Double Logistic Model fitted in the First Differences under Hypothesis H_A — Quarterly Data, 1981Q3–1995Q4**

Parameters of displaced logistic for the saturation share of imports as a function of Australian competitiveness

	1st quarter	*2nd quarter*	*3rd quarter*	*4th quarter*	*Price Sensitivity* γ
Location parameter	α_1	α_2	α_3	α_4	3.428
	0.1540	0.2120	0.0838	0.1355	
Minimum market share for imports:	θ_1	θ_2	θ_3	θ_4	
	0	0.0096	0.0191	0.0003	

Parameters of logistic trend in imports' market share		Error correction parameter in eqn (3.11)	Long- and short-run Armington elasticities (a)	Value of objective function (b)	Durbin-Watson (c)	Quasi-R^2 (levels) (d)	Quasi-R^2 (differences) (e)	Quasi-R^2 (differences) (f)
a	b	ψ	σ_{LR} at sample end	ϕ_A	DW(1)	$QR^2(W)$		$QR^2(\Delta \ln L)$
			3.428	0.3640	2.369		$QR^2(\Delta W)$	
17.696	-0.00609	2.5×10^{-7}	typical σ_{SR} 0.567	[ϕ_B = 1.178×10^{-3}]	DW(4) 1.331	0.835	0.706	0.923

* These are not the restricted maximum likelihood estimates; see text.

(a) σ_{LR} is calculated using (2.18); σ_{SR} is the sample mean of values calculated using (2.17).

(b) ϕ_A is the residual sum of squares from the finally fitted equation (3.9).

(c) DW(1) and DW(4) are the 1st and 4th-order Durbin-Watson statistics; they refer to residuals from the fitted equation (3.9). Corresponding statistics for (3.10) are 0.708 and 1.604.

(d) $QR^2(W)$ is the square of the simple correlation coefficient between $\hat{W}_t$ and W_t.

(e) $QR^2(\Delta W)$ is the square of the simple correlation coefficient between $\Delta\hat{W}_t$ and ΔW_t.

(f) $QR^2(\Delta \ln L)$ is the square of the simple correlation coefficient between $\Delta \ln \hat{L}_t$ and $\Delta \ln L_t$ from fitted equation (3.19).

Table 6.2; the corresponding fitted share values are shown in Figure 6.5. It will be seen that the change in starting values (in particular, setting γ to zero throughout) finds a parameter set yielding a lower ϕ_A (and therefore a higher likelihood) than those reported in Table 6.1. Using the solution of Table 6.2 as starting values, and searching with γ unconstrained, yields another solution almost indistinguishable from Table 6.2 with γ at solution returning to a number which is zero to four decimal places. As this provided the lowest ϕ_A among

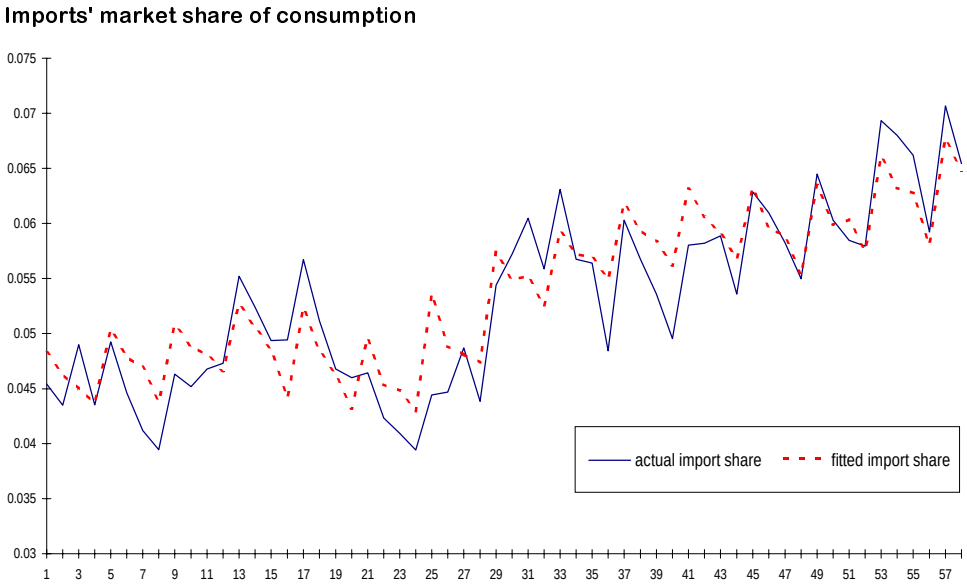

Figure 6.3 Projected and actual shares of imports in consumption, 1981Q3–1995Q4 (fitted values from Table 6.1 estimates)

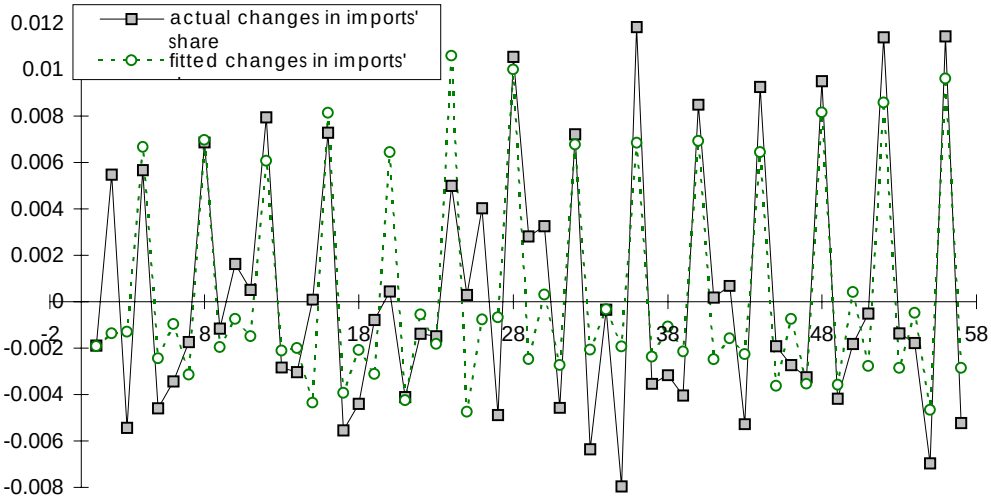

Figure 6.4 Projected and actual changes in shares of imports in consumption, 1981Q3—1995Q4 (fitted values from Table 6.1 estimates)

searches from several starting points, Table 6.2 is provisionally regarded as providing the restricted maximum likelihood estimates.

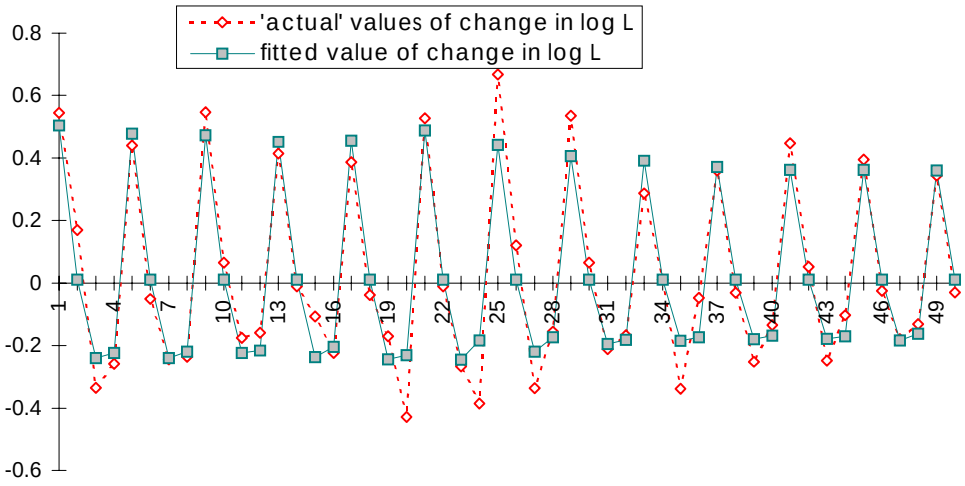

Figure 6.5 Fit of logarithmic changes in the log-ratio transform, Δ ln L, 1981Q3–1995Q4 (fitted values from Table 6.1 estimates)

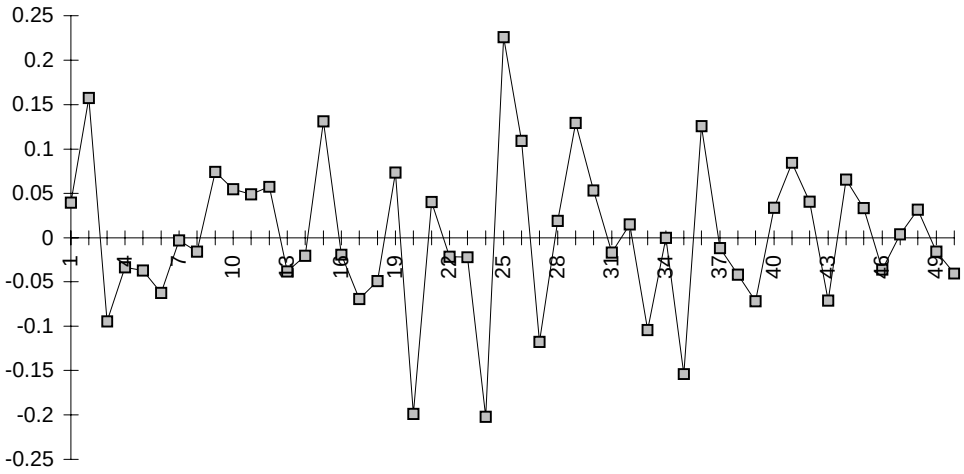

Figure 6.6 Residuals from fitted equation (3.9), 1981Q3—1995Q4 (fitted values from Table 6.1 estimates)

Table 6.2

Estimates of the Double Logistic Model fitted in the First Differences under Hypothesis H_A with Price Sensitivity set to Zero Quarterly Data, 1981Q3–1995Q4*

Parameters of displaced logistic for the saturation share of imports as a function of Australian competitiveness

	1st quarter	*2nd quarter*	*3rd quarter*	*4th quarter*	*Price Sensitivity* γ
Location parameter	α_1	α_2	α_3	α_4	**0** [by construction]
	0.1605	0.2519	0.0775	0.1409	
Minimum market share for imports:	θ_1	θ_2	θ_3	θ_4	
	0**	0.0096	0.0182	0.0003	

Parameters of logistic trend in imports' market share		Error correction parameter in eqn (3.11)	Long- and short-run Armington elasticities (a)	Value of objective function (b)	Durbin-Watson (c)	Quasi-R^2 (levels) (d)	Quasi-R^2 (differences) (e)	Quasi-R^2 (differences) (f)
a	b	ψ	σ_{LR} at sample end	ϕ_A	DW(1)	QR^2(W)		QR^2(Δ ln L)
			0	0.2841	2.02		QR^2(ΔW)	
19.465	-0.00713	2.5×10^{-7}	typical σ_{SR}	[ϕ_B = 1.58×10^{-3}]	DW(4)	0.763	0.759	0.924
			0		1.690			

* γ constrained to zero. To obtain the results reported here, the tolerances on the unbiasedness constraints had to be loosened somewhat, so that the intercept in the regression of W_t on $\hat{W}_t$ = -0.001 and the associated regression slope is 0.9998. When the solution was recomputed with γ unconstrained, the algorithm found a solution indistinguishable from the one reported in the table above.

** Corner solution.

(a) σ_{LR} is calculated using (2.18); σ_{SR} is the sample mean of values calculated using (2.17).

(b) ϕ_A is the residual sum of squares from the finally fitted equation (3.9).

(c) DW(1) and DW(4) are the 1st and 4th-order Durbin-Watson statistics; they refer to residuals from the fitted equation (3.9). Corresponding statistics for (3.10) are 0.392 and 1.621.

(d) QR^2(W) is the square of the simple correlation coefficient between $\hat{W}_t$ and W_t.

(e) QR^2(ΔW) is the square of the simple correlation coefficient between $\Delta\hat{W}_t$ and ΔW_t.

(f) QR^2(Δ ln L) is the square of the simple correlation coefficient between $\Delta \ln \hat{L}_t$ and $\Delta \ln L_t$ from fitted equation (3.19).

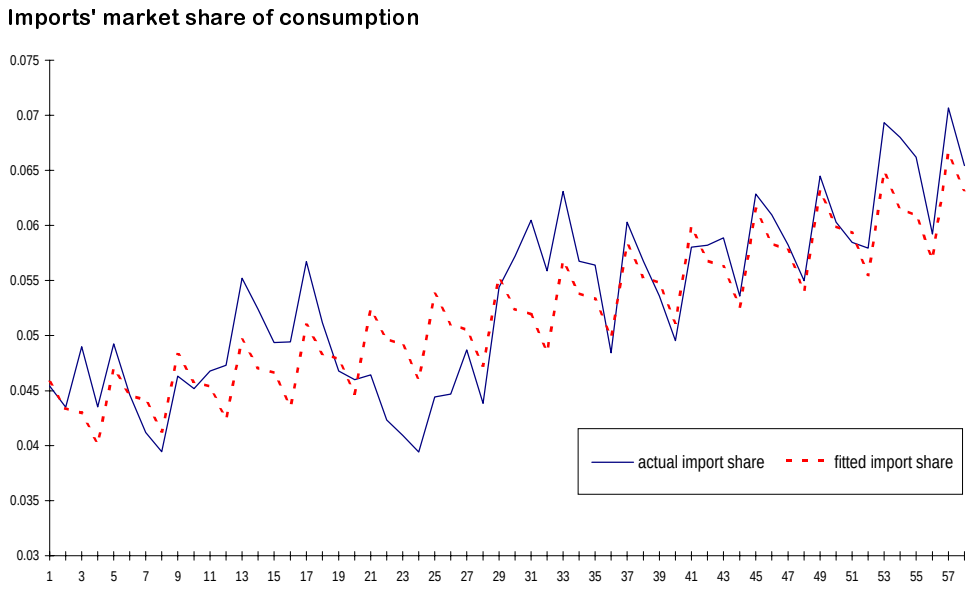

Figure 6.7 Projected and actual shares of imports in consumption, 1981Q3–1995Q4, when price influences are removed (i.e., $\gamma = 0$) (estimated under H_A). This corresponds to the provisional restricted maximum likelihood estimates of the parameters.

Figure 6.2 reveals that most of the variation in relative prices took place towards the middle of the period spanned by the old data. The additional observations added by the new data include relatively little variation in competitiveness. Thus in the new data, accounting for seasonality and pure trend becomes relatively more important for obtaining a good fit and for maximizing likelihood. Consequently, trend and seasonal components alone give a good account in Figure 6.7 of the variation in imports' market share, especially over the latter period of the sample. In terms of the likelihood, the failure to track the drop between quarters 18 and 28 is more than compensated for by the good fit obtained between quarters 36 and 58.

6.1 The bad and the good news revisited

In this subsection the estimates from Table 6.1 are taken as more plausible than those from Table 6.2. The strength of the trend against the local product according to the new data is somewhat stronger than that suggested by the old ($b = -0.00609$ versus -0.00525, according to Tables 6.1 and 5.1); but the effect of competitiveness in changing market shares is lower ($\gamma = 3.4$ versus 4.8).

Figure 6.8 is Figure 5.7 redrawn on the basis of Table 6.1. With the new data suggesting a considerably lower long-run elasticity of substitution between domestic and foreign consumption goods, the scope for lowering the ultimate level at which the market becomes saturated with imports is a good deal narrower in Figure 6.8 than in Figure 5.7.

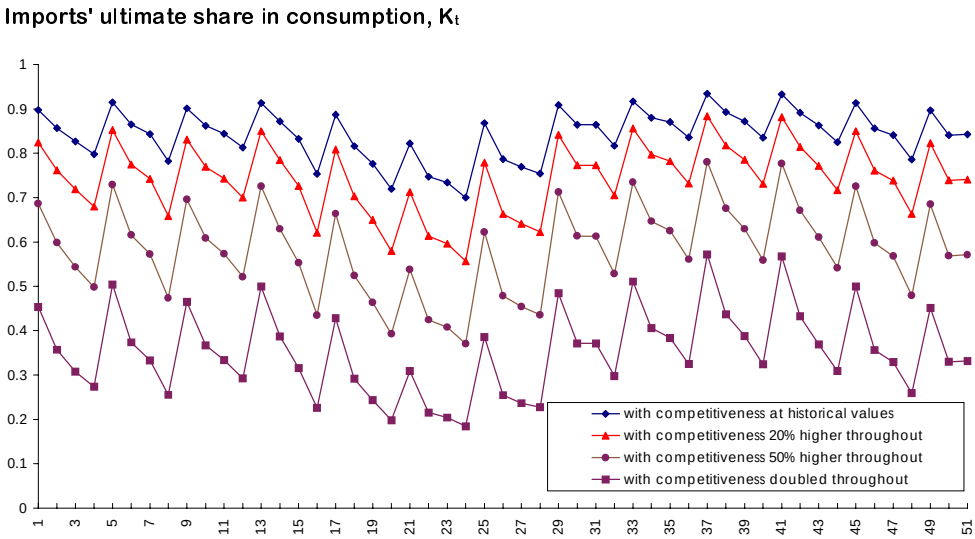

Figure 6.8 Estimated values of imports' ultimate share of consumption at levels of competitiveness that obtained during the sample, 1981Q3–1995Q4 (new data, Table 6.1 estimates)

Likewise, the new data suggest a typical short-run elasticity of substitution that is a good deal lower than that suggested by the old data (0.57 versus 0.79 — see Tables 5.1 and 6.1). A loss in competitiveness therefore is projected to do less harm according to the picture projected by Table 6.1; on the other hand, the structural shift in favour of imports is stronger.

These ideas become firmer if we project the market share of imports in consumption using the parameter values of Tables 5.1 and 6.1. In Figures 6.9 and 6.10, market share is projected over the period 1996Q1 through 2042Q4. Three sets of assumptions are used:

- competitiveness remains as it was in 1995Q4 — (in Figure 6.9);
- competitiveness improves at 4 per cent per annum
- competitiveness deteriorates at 4 per cent per annum

} (in Figure 6.10).

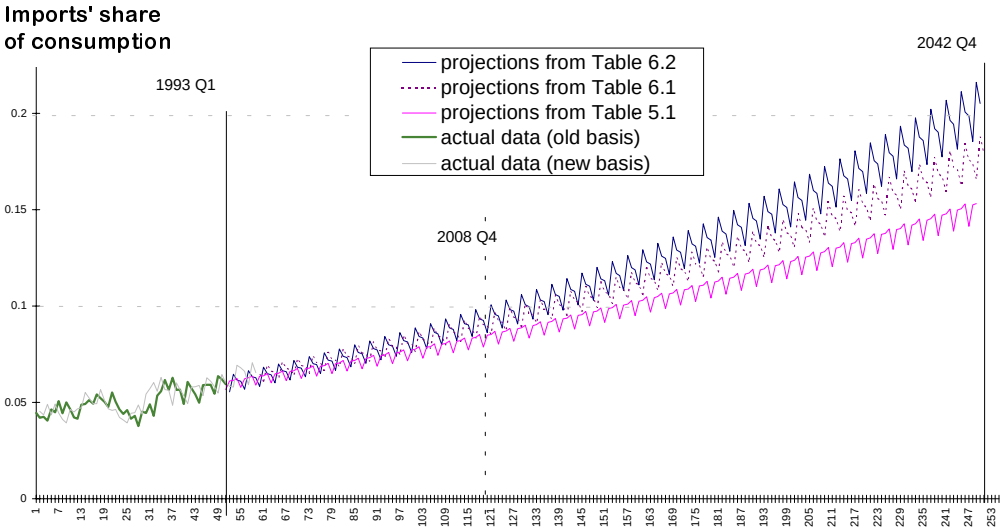

Figure 6.9 Projected trajectories of imports' share in consumption according to estimated parameters shown in Tables 5.1, 6.1 and 6.2 when Australian competitiveness remains at its 1995Q4 value. The stronger trend in the new data is evident. With $\gamma = 0$, or with P_t fixed, the ceiling parameter K_t is subject to a stationary seasonal pattern (see eqn (2.4)); the gaps between the seasonal low and the seasonal high in the middle trajectory increase to almost 13 percentage points as $t \to \infty$. This corresponds to a seaonal high of K equal to 0.929 and a seasonal low equal to 0.801.

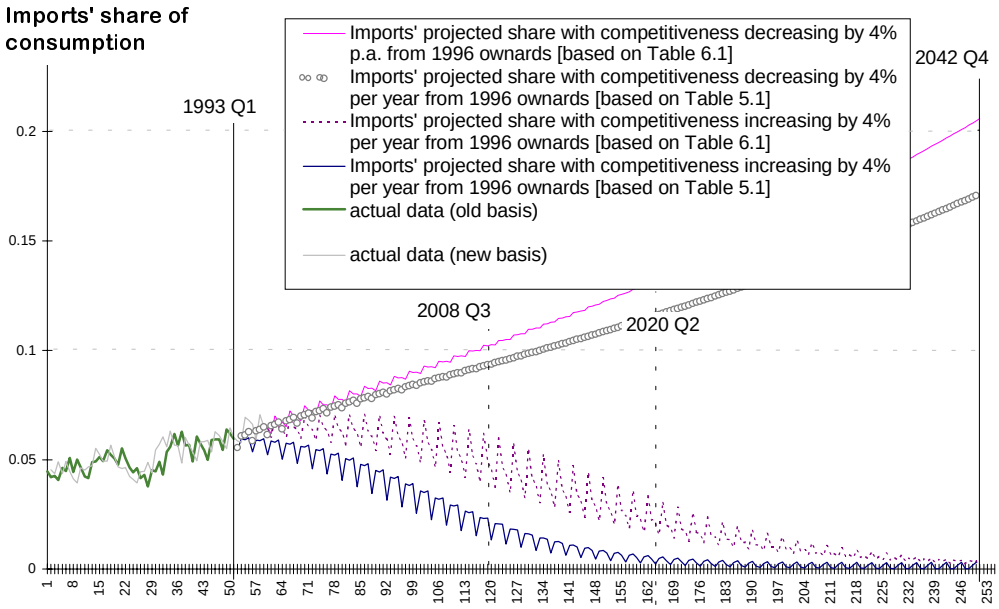

Figure 6.10 Projected trajectories of imports' share in consumption according to estimated parameters shown in Tables 5.1 and 6.1 when Australian competitiveness increases or deteriorates by 4 per cent per annum

Initially unexpected results are the seemingly explosive nature of the seasonal pattern in Figure 6.9 and, in Figure 6.10, the 'dinosaur skeleton' pattern displayed in the second-lowest trajectory. In the latter case, the already prominent seasonality evident in the sample is amplified in the middle of the projection period. The

explanation for the pattern of seasonal behaviour evident in Figure 6.9 is given in the caption to that figure.

Recall that the procedure of removing seasonality was developed in explorations with the old data. As we have seen above, the method works well within-sample for both data sets. Although the lowest curve in Figure 6.10 (which is based on coefficients estimated from the old data) does show a similar pattern of amplified seasonality during the middle part of the adjustment path, the effect is less pronounced. At all events, seasonality is a minor issue for the out-of-sample projections; if desired, a milder form of seasonality can be imposed on the annualized projections shown in Figure 6.9.

6.2 Summary of estimation results

Point estimates of the long-run elasticity of substitution between domestic and imported consumer goods are sensitive to the data set used — the (old) data terminating in the 1st quarter of 1993 yield a point estimate of 4.8, whereas

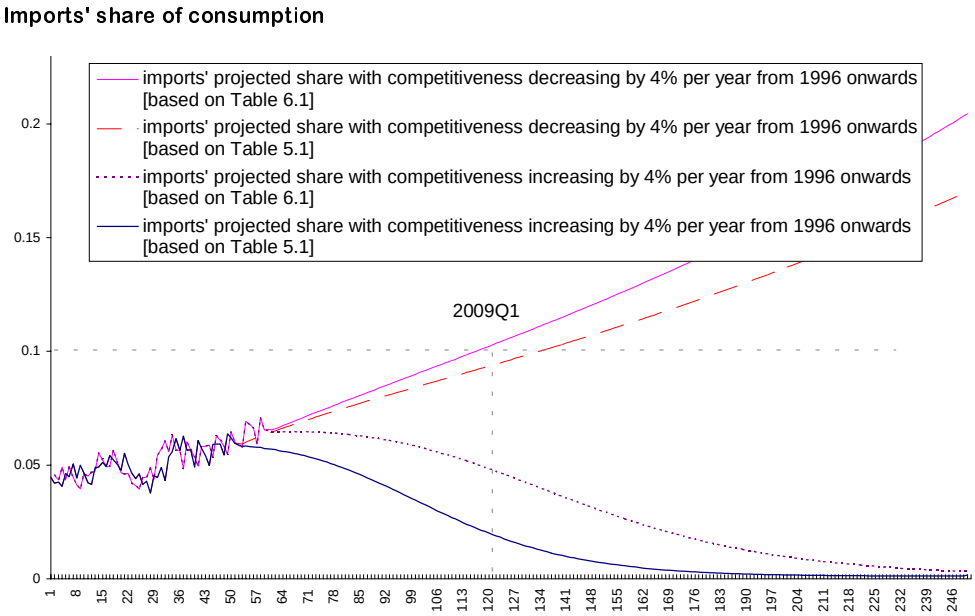

Figure 6.11 Projected trajectories of imports' share in consumption according to estimated parameters shown in Tables 5.1 and 6.1 when Australian competitiveness increases or deteriorates by 4 per cent per annum, with seasonal variation removed in projection period

the rebased (new) data terminating in the 4th quarter of 1995 suggest a value of about 3.4, or alternatively, zero. Estimates of b corresponding to the old and the new data sets range between -0.005 and -0.007. These values for the pure trend parameter are relatively insensitive to variations in model specification (such as the elimination of the influence of prices) and to minor variations in data handling, suggesting that it may have a relatively tight sampling distribution.

We can, however, be less sure about the accuracy of the point estimate of price responsiveness (γ), although contrasting Figures 5.2 and 5.6 provides rather strong evidence in favour of the significant effect of competitiveness on market share. In the case of the new data, unfortunately, the comparison between

Figures 6.3 and 6.7 is far from convincing: the latter figure reveals that seasonality and pure trend dominate the final four years of the data series on imports' market share, which is enough to tip the balance against finding any relative price effect. A relative maximum of the likelihood function yields the γ estimate of 3.4 underlying Figure 6.3.

Point estimates of short-run (quarterly) elasticities of substitution σ_{SR} are variable in the model, responding both to changed competitiveness at a given point of time, and to time at a given level of competitiveness. Estimated σ_{SR} have typical values of about 0.8 and either zero or 0.6 in the old and the new data sets respectively.

7. Concluding remarks

7.1 Scope for further work

It is disappointing not to be able to report standard errors. The most likely avenues for generating them are either (a) the use of a numerical Hessian of the likelihood function computed in GAUSS or a similar package; or (b) the use of a bootstrapping procedure to generate the sampling distributions of the estimators empirically. At this stage it is not clear whether the mixture of formal and informal procedures used in Sections 5 and 6 can be automated sufficiently to make the latter project feasible. Further, both approaches are biased towards model confirmation since the specification was generated interactively with the data.

It is clear that modelling price competition exclusively through the seasonal asymptotes is only one alternative; with an abundant data supply it would be possible to model both the slope and location parameters of the time trend. Given the current limited supply of data, however, it is difficult to identify feasible extensions of the model. One option might be to abandon the estimation of the seasonal influences altogether, and to work with seasonally adjusted data. Despite the obvious loss of control in the specification of the fundamentals of the market adjustment process and the attendant difficulties of interpretation, the removal of six parameters is attractive.

7.2 Prospects for market penetration by imports

In response to the issue raised in the *Introduction*, it seems that although we are indeed in phase A of Figure 1.1, the potential ultimate share of imports is high, possibly close to 100 per cent, and at least in the 80s at the level of competitiveness that prevailed in late 1995. (But keep in mind that market saturation is estimated to take of the order of half a century, and we should be very cautious about taking our forecasting abilities this seriously.)

Since it is always hazardous to project trends any distance into the future, perhaps we might take somewhat seriously the implications of the above results for just the next decade.

The rate of structural change in the market for consumption goods seems to be moderate. On the most pessimistic premise canvassed here (a decline in competitiveness of 4 per cent per year), imports' (average annual) share of the market does not reach 10 per cent until 2009. With no change in competitiveness, the 10 per cent mark is reached about 6 years later (according to the Table 6.1 estimates).

There is scope to limit this penetration by lowering the cost of locally produced consumer goods. If the estimated price responsiveness implied by Table 6.1 is believed, then a maintained improvement of competitiveness of about $2\frac{1}{2}$ per cent per year would arrest the decline in local market share for at least the next 10 years. If the estimates in Table 5.1 are believed, the sustained improvement in competitiveness need only be about $1\frac{1}{2}$ per cent per annum. To these statements the caution must be added that the measures of price responsiveness on which they depend are heavily influenced by the market experience of 1986-1988 (the 'dip' in Figure 5.2).

This was a period of relatively great variation in competitiveness (the 'peak' in Figure 6.2). Largely driven by the real exchange rate, there were highly collinear movements in the prices of the different components of aggregate imports, which helps to justify the invocation of Hicks' composite commodity theorem to allow the treatment of imported consumer goods as a single item. On the other hand, when changes in the relative prices of the domestic and the imported aggregates involve significant price variation within the aggregates, the applicability of the estimates derived above is on less firm ground.

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