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INTERTEMPORAL SUBSTITUTION IN IMPORT DEMAND AND HABIT FORMATION

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SUMMARY

To study non-durable import demand, we extend previous work done by Clarida (1994) and Ceglowski (1991) by considering a two-good version of the lifecycle model in which we introduce time-non-separability in the households' preferences. The model is estimated using quarterly data for the USA and France. Using the information contained in the observed stochastic and deterministic trends, we derive a cointegration restriction used to estimate curvature parameters of the instantaneous utility function. The remaining parameters are estimated in a second step by GMM. The constancy of the different parameters is investigated, in both the long and the short run. Habit formation turns out to be an important factor of import demand. © 1998 John Wiley & Sons, Ltd.

1. INTRODUCTION

A large number of empirical applications on trade balance and imports are now explicitly drawn from the optimality conditions of an intertemporal maximization programme, under the assumption of rational expectations. These models capture explicitly a simple idea that is missing in earlier work albeit central to macroeconomics: there is an ever-present competition of resources between today and future periods. As far as imports are concerned, such models allow us to consider, along with the relative price of imports, the real interest rates as a second channel through which policy could affect the trade balance (Ceglowski, 1991; Amano and Wirjanto, 1996).[1]

Taking interest rates so explicitly into account is in sharp contrast to most of the existing empirical literature on import demand which is mainly 'atheoretical' in the sense that the so-called import demand models are derived from a pure empirical exercise (see, for example, Urbain, 1992 and the references therein). Looking back to the existing literature dealing with the empirical modelling of trade flows, and more precisely of import demand, one can roughly derive two general classes of studies.

The first part, and certainly the most important, of the applied econometrics literature dealing with import demand models has been based on what has often been called pseudo-reduced-form models (see the survey of Goldstein and Khan, 1985) or empirically derived dynamic specifications (see, *inter alia*, Urbain, 1992; Asseery and Peel, 1991; Mah, 1994). While these models

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[1] A related question is to know to what extent the very high interest rates experienced by Europe since the beginning of the 1980s could be held responsible for the existing surplus of current accounts.

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usually have 'naive' theoretical foundations based on some form of the imperfect substitutes model surveyed by Goldstein and Khan (1985), they nevertheless have the advantage of enabling the time-series properties of the data to be taken fully into account so that the resulting models (if derived within a coherent modelling framework) are often statistically well specified (see Urbain, 1995).

The second class of studies is implicitly motivated by the Lucas (1976) critique which argues that the parameters of traditional macroeconometric models depend crucially on parameters governing the processes used to form agents' expectations and are unlikely to remain stable in a changing economic environment. As a response to this, a number of papers have focused on the estimation of theoretical intertemporal optimization models with rational expectations assumed to have an explicit and direct structural interpretation (see, *inter alia*, Ceglowski, 1991; Clarida, 1994; Kollintzas and Husted, 1984; Husted and Kollintzas, 1987; Gagnon, 1988). Although these models have some strong theoretical motivation, it must be pointed out that the statistical properties of the time-series data used are often not taken into account,[2] or at least not fully exploited. One of the interesting aspects of working in this set-up is that a number of other factors affecting trade balance, such as real interest rates changes, emerge from the analysis.

This paper integrates the time-series characteristics of the data into the study of a theoretically based dynamic model for consumer non-durable imports. If the driving forces of the economy are indeed non-stationary processes, as is now almost well accepted in the literature, then intertemporal optimization models lead in general to two types of testable restrictions (see, for example, Canova, Finn, and Pagan, 1994). First, there are long-run restrictions reflecting the fact that there are generally more variables to be modelled than there are independent forcing processes. Second, there are restrictions upon the (short-run) dynamics of the system.

In general, long-run restrictions are not rejected by the data. For instance, with strongly separable addilog preferences, the intertemporal import demand theory predicts that the log of demand for import goods, the log of demand for domestic goods, and the log of the relative price of imports, if I(1) processes, are cointegrated. This long-run relation describes the potential substitution between current imports and current consumption of domestic goods. Clarida (1994) finds strong support in the data for such cointegration vectors. As stressed by Ogaki (1992), the interest of the approach is that the estimated parameters of the cointegration vector are sufficient to identify the curvature parameters of the utility function (i.e. the intertemporal elasticity of substitution for the two goods). It appears that the intertemporal elasticity of substitution for consumer imports is significantly larger than the one for domestic consumption.[3]

Turning our attention to the short-run dynamics, the evidence in favour of the theory is less clear (see Hansen and Singleton, 1982 for consumption, Ceglowski, 1991 for imports and Otto, 1992 for the current account) or not investigated at all. Moreover, the constancy of the short-run parameters is very seldomly analysed. These are important issues since these short-run relations describe the substitution of consumption between two points in time, which is the interesting mechanism of present-value models. Empirical rejection of the validity of these can cast some doubts on the validity of the underlying theoretical model.[4]

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[2] An exception being the paper by Clarida (1994).

[3] As stressed later, this interpretation of the curvature parameters as the intertemporal elasticity of substitution is no longer valid when habit formation is introduced.

[4] Effectively, long-run restrictions, such as cointegration restrictions, are not sufficient to provide strong empirical support to a given theoretical model since different theoretical models can lead to similar long-run relations among a given set of variables.

The issues that we explore in this paper are threefold:

(1) The existing US-based empirical analysis of the stationarity restrictions implied by the present-value model has led to the conclusion that the intertemporal elasticity of substitution for non-durables imports is three times the one for domestic goods. We would like to investigate whether this important result is confirmed by French data and whether this long-run information contained in the data can be used to improve our knowledge of the dynamics.
(2) We generalize the set-up used in Ceglowski (1991), Clarida (1994), and Amano and Wirjanto (1996) in order to allow for a richer dynamics. In particular, we investigate whether introducing intertemporal non-separability in households' utility could be helpful to account for imports dynamics in the face of changes in interest rates. Our intuition is that the pure forward-looking dynamics generated by the basic present-value model are too restrictive and that the introduction of time-non-separability introduces richer dynamic structures with some backward-looking elements.
(3) Although often not explicitly investigated, we pay some attention to the empirical success (or failure) of our theoretical specification by investigating the constancy of the deep structural parameters estimates (in the sense of Lucas).

The rationality for introducing time-non-separable preferences is to be found in the works of socio-psychologists (see Argyle, 1987) and biologists (Helson, 1964). Habit formation is one form of time-non-separability which has been extensively studied in consumption theory. The idea dates back to Duesenberry (1949). It amounts basically to assuming that tastes are changing and that these changes depend on past decisions (i.e. past consumption levels or expenditures level). If we consider only the recent contributions, Muellbauer (1988), Eichenbaum, Hansen, and Singleton (1988), Ferson and Constantinides (1991), and Ogaki and Park (1994) find that habit formation helps to account for consumption dynamics. However, to our knowledge, habit formation has never been introduced in studies of import demand or current account determination. Our purpose is to investigate, in the case of imports, whether 'ignoring habits or other forms of non-separability may explain the frequent rejection of the life cycle hypothesis' (Winder and Palm, 1996).

The structure of this paper is as follows. In Section 2 we present the theoretical model, derive the intertemporal substitution properties, and point out how one can derive a cointegration restriction from this set-up. Section 3 presents the data series and some univariate time-series properties. Section 4 reports the estimation results and investigates the issue of the constancy of the parameters. A final section offers conclusions.

2. THE MODEL

The representative household has preferences defined over the services q_m and q_d provided by the acquisitions of non-durable imported consumption goods c_m and non-durable domestic consumption goods c_d. The purchase of goods is transformed into services according to the following relation:

$$q_{it} = \frac{c_{it} - \gamma_i c_{it-1}}{1 - \gamma_i} \qquad i = m, d \tag{1}$$

The parameters $\gamma_m, \gamma_d \in]-1, 1[$ measure the extent to which past purchases affect current services. If they are positive, the household is subject to habit formation: the more the consumer has purchased in the previous period, the more he or she has to purchase in order to attain the same level of satisfaction. A negative γ_i indicates that the good presents some durability (see e.g. Ferson and Constantinides, 1991): in that case, the accumulation of past flows of purchases contributes to satisfaction. This accumulation is, in our case, limited to one period. Note that, even with non-durable goods, one cannot fully exclude the presence of the effect of some inventory holdings. If these parameters are zero, the utility function is time-separable as in Clarida (1994) and Ceglowski (1991). The specific form of the instantaneous utility function of the representative household is:

$$U(q_{dt}, q_{mt}) = \sum_{i=m,d} \frac{\exp(\mu_i t)}{1-\alpha_i} q_{it}^{1-\alpha_i} \tag{2}$$

which is a special case of the form used by Eichenbaum, Hansen, and Singleton (1988) with no utility for leisure.[5] The parameters $\alpha_i > 0$ are called curvature parameters. Their inverse can be interpreted as intertemporal elasticities of substitution (IES) when there is no habit formation. In the presence of habits, the IES has a more complicated expression which is derived in Boldrin, Christiano and Fisher (1995). The trend terms μ_i could be interpreted as deterministic technological progress in the transformation of purchases into services. Let us denote the stock of assets and the labour income of the representative household respectively by a_t and y_t. Labour income is assumed to be exogenous. The representative household selects $\{c_{ds}, c_{ms}, a_s; s \geqslant t\}$ in order to:

$$\max \mathrm{E}\left[\sum_{s=t}^{\infty} \theta^{s-t} U(c_{ds}, c_{ds-1}, c_{ms}, c_{ms-1}) \mid \Omega_t\right] \qquad 0 < \theta < 1$$

subject to

$$\sum_{i=m,d} p_{is} c_{is} + a_s = (1+r_s) a_{s-1} + y_s \text{ for all } s \in [t, \infty]$$

$$a_{t-1}, c_{mt-1}, c_{dt-1} \text{ given}$$

where p_{is} is the price of good i, Ω_t the information set at time t, r_t the rate of return on assets, and θ the rate of time preference (i.e. the subjective discount rate). To ensure that the intertemporal objective function is finite and that the intertemporal problem is well defined, we assume that $\theta \exp(\mu_i) < 1$ for $i = m, d$. The optimal choice of consumption, import and savings at time t should satisfy:

$$p_{it}\lambda_t = \exp(\mu_i t)\, q_{it}^{-\alpha_i} - \gamma_i \theta \exp(\mu_i (t+1))\, \mathrm{E}_t[q_{it+1}^{-\alpha_i}] \qquad i = m, d \tag{3}$$

$$\lambda_t = \theta \mathrm{E}_t[(1+r_{t+1})\, \lambda_{t+1}] \tag{4}$$

$$\mathrm{E}_t\left[\lim_{i\to\infty} \frac{a_{t+i}}{\prod_{s=t}^{i}(1+r_s)}\right] = 0 \tag{5}$$

[5] The utility function (2) displays strong separability between the two goods. Ogaki (1992) points out that the long-run restriction implied by the maximisation of (2) is still valid under a concave transformation of (2). This, of course, is not the case for the short-run restrictions.

λ denotes the Lagrange multiplier associated to the accumulation constraint and $E_t[\cdot] = E[\cdot \mid \Omega_t]$. Equation (5) excludes Ponzi games in which the value of the consumer's debt increases in the limit more rapidly than the compound interest rate.

In this framework, today's consumption of imports can be substituted for future consumption (intertemporal substitution) or for today's consumption of domestic goods (contemporaneous substitution).

2.1. Intertemporal Substitution

From equations (3) and (4), we get:

$$q_{it}^{-\alpha_i} = E_t\left[\gamma_i \theta \exp(\mu_i) q_{it-1}^{-\alpha_i} + \frac{\theta \exp(\mu_i)}{R_{it+1}}(q_{it+1}^{-\alpha_i} - \gamma_i \theta \exp(\mu_i) q_{it+2}^{-\alpha_i})\right] \tag{6}$$

where

$$R_{it+1} = \frac{1}{(1 + r_{t+1})} \frac{p_{it+1}}{p_{it}} \tag{7}$$

for $i = m, d$. R_i are the commodity-specific real discount factors. They represent the opportunity cost of postponing import services (resp. domestic good services) in period t in order to increase import services (resp. domestic good services) in period $t + 1$. Equation (6) implies that an increase in the interest rate should induce households to substitute future consumption for current consumption, as long as the parameters α_i are positive. If $\gamma_i = 0$, equations (1) and (6) imply

$$E_t\left[\frac{\theta \exp(\mu_i)}{R_{it-1}}\left(\frac{c_{it+1}}{c_{it}}\right)^{-\alpha_i} - 1\right] = 0$$

An increase in the real interest rate by one per cent induces a rise in planned consumption of good i by $1/\alpha_i$ per cent. To be more precise, this holds only if foresights are perfect. However, as stressed by McLaughlin (1995), $1/\alpha_i$ should produce high-quality approximations to the true effects, even with rational expectations and in the absence of consumption insurance. Notice that, following Kim (1993), when within-period preferences are additively separable as in equation (2), changes in the interest rate and in the commodity price have the same intertemporal substitution effect on a commodity demand.

When habit formation is allowed for, the effect of a rise in interest rate is more complex, since agents recognize the impact of today's choices on their future tastes. In that case, Constantinides (1990) shows that there is a gap between the relative risk aversion and the intertemporal elasticity of substitution. In that case, $1/\alpha_i$ is an invalid measure of the intertemporal substitution effect. Nevertheless, the IES can still be recovered in the presence of habit formation as shown by Boldrin, Christiano, and Fisher (1995) in a deterministic framework. They demonstrate the difference between risk aversion and the IES in the presence of habit formation of a more general type than the one used here. Adapting their derivation to our case, the IES for good i is given by:

$$\frac{1}{\alpha_i}(1 - \gamma_i) \frac{1 - \gamma_i \theta \exp(\mu_i)/(1 + g_i)^{\alpha_i}}{1 + \gamma_i^2 \theta \exp(\mu_i)/(1 + g_i)^{\alpha_i + 1}} \tag{8}$$

where g_i is the average of $\Delta \ln c_{it}$ for $i = m, d$. Notice that this expression is a correct measure of the IES only in a deterministic context and that it is subject to the same criticism by McLaughlin (1995) as in the case without habit formation.

2.2. Contemporaneous Substitution

Combining the two equations implied by (3), the relative price between the two goods should equal the marginal rate of substitution of these goods:

$$\frac{p_{dt}}{p_{mt}} = \exp((\mu_d - \mu_m)t)\ \frac{q_{dt}^{-\alpha_d} - \gamma_d \theta \exp(\mu_d) \mathrm{E}_t[q_{dt+1}^{-\alpha_d}]}{q_{mt}^{-\alpha_m} - \gamma_m \theta \exp(\mu_m) \mathrm{E}_t[q_{mt+1}^{-\alpha_m}]} \tag{9}$$

dividing both sides by $q_{dt}^{-\alpha_d} q_{mt}^{-\alpha_m}$ and taking logs:

$$\begin{aligned} &\alpha_m \ln q_{mt} - \alpha_d \ln q_{dt} - \ln\left(\frac{p_{dt}}{p_{mt}}\right) + (\mu_d - \mu_m)t \\ &= \ln\left\{1 - \gamma_m \theta \exp(\mu_m) \mathrm{E}_t\left[\left(\frac{q_{mt}}{q_{mt+1}}\right)^{\alpha_m}\right]\right\} - \ln\left\{1 - \gamma_d \theta \exp(\mu_d) \mathrm{E}_t\left[\left(\frac{q_{dt}}{q_{dt+1}}\right)^{\alpha_d}\right]\right\} \end{aligned} \tag{10}$$

Following the arguments of Ogaki and Park (1994), it can be shown that equation (10) implies that $\ln c_{mt}$, $\ln(p_{dt}/p_{mt})$ and $\ln c_{dt}$ be cointegrated, as long as these variables are integrated processes of order one. This can be intuitively shown adapting the approximation of the logarithm of equation (1) proposed by Muellbauer (1988):

$$\ln q_{it} \approx \mathrm{cst} + \ln c_{it} + \left(\frac{1 - g_i}{\frac{1-\gamma_i}{\gamma_i} + g_i}\right) \Delta \ln c_{it} \tag{11}$$

where g_i is the average of $\Delta \ln c_{it}$. Using equation (11) in (10):

$$\begin{aligned} &\alpha_m \ln c_{mt} - \alpha_d \ln c_{dt} - \ln\left(\frac{p_{dt}}{p_{mt}}\right) + (\mu_d - \mu_m)t \approx \mathrm{cst} \\ &+ \ln\left\{1 - \gamma_m \theta \exp(\mu_m) \mathrm{E}_t\left[\left(\frac{q_{mt}}{q_{mt+1}}\right)^{\alpha_m}\right]\right\} - \alpha_m \left(\frac{1 - g_m}{\frac{1-\gamma_m}{\gamma_m} + g_m}\right) \Delta \ln c_{mt} \\ &- \ln\left\{1 - \gamma_d \theta \exp(\mu_d) \mathrm{E}_t\left[\left(\frac{q_{dt}}{q_{dt+1}}\right)^{\alpha_d}\right]\right\} + \alpha_d \left(\frac{1 - g_d}{\frac{1-\gamma_d}{\gamma_d} + g_d}\right) \Delta \ln c_{dt} \end{aligned} \tag{12}$$

which provides us with a stationary/cointegration restriction, since the right-hand side of equation (12) is expressed in terms of covariance-stationary variables.

When $\gamma_i = 0 \; \forall \; i$ as in Clarida (1994), using equation (3) and taking logs leads to

$$\alpha_m \ln c_{mt} - \alpha_d \ln c_{dt} - \ln\left(\frac{p_{dt}}{p_{mt}}\right) + (\mu_d - \mu_m)t = 0 \tag{13}$$

which requires $\ln c_{mt}$, $\ln(p_{dt}/p_{mt})$ and $\ln c_{dt}$ be cointegrated.[6] Following Clarida (1994), the two remaining common stochastic trends among the three I(1) variables can be identified as the log of the marginal utility of wealth and as a permanent technological shock to the supply of imported goods and, henceforth, the relative prices. Clarida proposes also to compute the standard Marshallian price elasticity (at constant total expenditures) and the expenditure elasticity (at constant prices) of the demand for imports:

$$\eta_{c_m \cdot p_m/p_d} = -\frac{1}{\alpha_m}\left[1 - \frac{(1-\alpha_m)(1-s)}{(\alpha_m s/\alpha_d) + (1-s)}\right]$$

$$\eta_{c_m \cdot (c_d + p_m c_m/p_d)} = \frac{\alpha_d}{\alpha_m}\left[\frac{1}{s + (\alpha_d/\alpha_m)(1-s)}\right]$$

where s is the share of spending that falls on domestic goods.

With $\gamma_i \neq 0$, $i = m, d$, we can no longer express the demand for goods as an explicit function of prices and of the marginal utility of wealth λ_t. The two remaining common stochastic trends can no longer be identified but are still related to the marginal utility of wealth and to the permanent technological shock to the supply of imported goods. We can also no longer compute the instantaneous Marshallian elasticities.

3. EMPIRICAL RESULTS

As surveyed by Hall (1993), Ogaki (1993a), and Fève and Langot (1995), non-linear dynamic rational expectations models such as the one underlying equations (3) and (4) are often estimated by some versions of Hansen's (1982) Generalized Methods of Moments which yields, under some regularity conditions, asymptotically normally distributed and consistent estimates of the parameters of interest. These regularity conditions, however, include a strict stationarity assumption which is rarely met in practice with macroeconomic data series. Notice that if the time series used in equations (3) and (4) are characterized by deterministic trend components, then asymptotic normality still holds (Andrews and McDermott, 1995). The case of unit root processes is different and is ruled out by these regularity conditions.[7] An empirical popular response to this issue is to 'detrend' the data by, for example, reformulating the model in terms of growth rates. The approach we follow in this paper is, however, different as we want to investigate whether the observed non-stationarity in the data may be helpful in estimating some of the parameters. In an analogy to the unit root/cointegration literature in linear models one may indeed expect that the imposition (instead of estimation) of the appropriate unit roots/cointegration restriction will avoid the GMM estimates to have non-standard limiting distribution and enable us to estimate some parameter super-consistently.

[6] In Clarida (1994), stationary preference shocks are introduced so that the right-hand side of equation (13) is not 0 but a function of these shocks.

[7] To our knowledge, no exact results seem to exist yet on the behaviour of GMM for non-linear Euler equations with (co)integrated processes.

Our empirical analysis will thus proceed in three major steps. In the first step, we use the cointegration/stationarity restrictions derived in equation (12) to estimate the curvature parameters of the instantaneous utility function. Given that these parameters are super-consistently estimated (see Phillips and Hansen, 1990), we fix these parameters at their estimated values and estimate in a second step the remaining parameters of equation (6) by the Generalized Methods of Moments (Hansen, 1982). The last step of the empirical analysis consists of investigation of the potential parameter (non)-constancy of our retained specification and estimation results.

The data we use in this paper are quarterly seasonally adjusted covering the period 70:01–94:01 for France (97 observations) and 67:01–94:03 for the USA (111 observations). French data are built on the basis of the quarterly national accounts (Comptes nationaux trimestriels (INSEE)). Non-durables include subsectors U02 (meat and milk products, other products from the food industry) and U06 (drugs, textiles, clothing, shoes, leather, furnitures, printing). For the USA, both the source and the construction of the data follow Ceglowski (1991) for an extended sample size. Following Shapiro (1986), the required rate of return is the short-run real return (three-month treasury bill rate for the USA and 'taux de l'argent au jour le jour des effets privés' for France) plus a constant risk premium of 2% per quarter.

Some summary statistics are reported in Table I. From this table it appears that the ratio of imports to domestically produced non-durables is smaller on average in the USA, which is not surprising. One noticeable difference between the two countries is the higher volatility (measured by the standard deviation) of imports growth rates in the USA. The commodity-specific discount factors are comparable across countries and the one of imports is always more volatile than the one of domestic goods.

Table I. Summary statistics

Country	Variable	Average	Std dev.	Minimum	Maximum
France	c_{mt}/c_{dt}	0·1405	0·0460	0·0641	0·2198
	p_{mt}/p_{dt}	0·9904	0·0902	0·7926	1·1846
	c_{mt}/c_{mt-1}	1·0179	0·0252	0·9378	1·0778
	c_{dt}/c_{dt-1}	1·0045	0·0090	0·9841	1·0313
	R_{mt}	1·0318	0·0252	0·9612	1·0938
	R_{dt}	1·0275	0·0090	1·0102	1·0510
USA	c_{mt}/c_{dt}	0·0589	0·0158	0·0332	0·0905
	p_{mt}/p_{dt}	1·0006	0·1318	0·7853	1·3087
	c_{mt}/c_{mt-1}	1·0151	0·0750	0·7665	1·3442
	c_{dt}/c_{dt-1}	1·0046	0·0078	0·9820	1·0342
	R_{mt}	1·0238	0·0283	0·9430	1·1077
	R_{dt}	1·0244	0·0073	1·0063	1·0045

3.1. Time Series Properties of the Data

The first step in the analysis is the computation of some standard univariate unit root tests in order to obtain empirical evidence in favour of or against the assumption of stochastic trends in our data. This step, although usually considered more as a descriptive one in many empirical analyses, is of relative importance for our purpose since the stationarity/cointegration restriction

Table II. Unit root tests

	Variable		Levels				1st diff.		
		det.	ADF	PP	SP	det.	ADF	PP	SP
France	$\ln p_{mt}/p_{dt}$	cst, trd	−3·13	−2·89	−3·51	cst	−5·01	−6·77	−6·31
		cst	−0·84	−0·92			−4·71	−6·70	
	$\ln c_{mt}$	cst, trd	−3·36	−2·53	−2·16	cst	−4·17	−8·03	−8·18
		cst	−2·40	−2·71			−2·69	−6·14	
	$\ln c_{dt}$	cst, trd	−2·20	−3·37	−1·77	cst	−3·62	−12·58	−12·05
		cst	−2·52	−4·03			−2·89	−10·45	
	R_{mt}	cst, trd	−4·95	−6·36	−5·77	cst	−6·62	−13·61	−7·87
		cst	−3·97	−5·94			−6·65	−13·71	
	R_{dt}	cst, trd	−4·40	−6·09	−5·54	cst	−5·73	−16·47	−10·84
		cst	−1·42	−3·13			−5·75	−16·59	
USA	$\ln p_{mt}/p_{dt}$	cst, trd	−1·73	−1·44	−1·76	cst	−4·31	−7·57	−6·75
		cst	−1·91	−1·62			−4·32	−7·59	
	$\ln c_{mt}$	cst, trd	−2·54	−3·15	−3·12	cst	−5·16	−14·58	−13·60
		cst	−0·35	−0·80			−4·62	−13·72	
	$\ln c_{dt}$	cst, trd	−2·53	−2·31	−2·77	cst	−4·59	−11·32	−12·04
		cst	−0·84	−1·19			−2·84	−9·00	
	R_{mt}	cst, trd	−3·74	−6·62	−6·55	cst	−6·49	−17·83	−8·58
		cst	−3·56	−6·49			−6·52	−17·93	
	R_{dt}	cst, trd	−2·20	−4·77	−4·12	cst	−5·99	−16·55	−7·85
		cst	−2·09	−4·29			−6·02	−16·65	
5% crit.	values	cst, trd	−3·44	−3·44	−3·60	cst	−2·89	−2·89	−3·60
		cst	−2·89	−2·89			−1·95	−1·95	

(12) hinges on the assumption that $\ln c_{mt}$, $\ln(p_{dt}/p_{mt})$ and $\ln c_{dt}$ are well described by unit root processes with possible drifts.

Notice that even if all our data series do not contain unit roots, but some of them are better modelled as trend stationary processes, then a similar stationarity/cointegration restriction can still be derived by introducing concepts such as cotrending and by distinguishing between stochastic and deterministic cointegration following the terminology of Ogaki and Park (1994). We say that difference stationary processes are deterministically cointegrated if the cointegrating vectors annihilate both the linear deterministic trend and the stochastic trend components. If only the stochastic trends are annihilated, we say that the series are stochastically cointegrated (see, for example, Ogaki, 1993b).

Table II presents the outcome of Dickey and Fuller (1979), Phillips and Perron (1988) tests as well as the Schmidt and Phillips (1992) test which has the advantage of being invariant to the specification of the deterministic components.[8] Critical values for DF and PP statistics are given in MacKinnon (1991) while for SP, these are reported in Schmidt and Phillips (1992). In all cases the test statistics are computed for two different specifications of the deterministic part ('det.' in the table): a constant and a linear time trend (cst, trd) and a constant term alone (cst). The results of these test statistics can be summarized as follows: $\ln c_{dt}$, $\ln c_{mt}$ and $\ln p_{mt}/p_{dt}$ seem to be well

[8] For both Phillips–Perron (PP) and Schmidt–Phillips (SP) we use the Newey and West (1987) estimator of the long-run variances with four lags. Notice that the results are rather insensitive to the number of lags actually used. The same lag length is used for the Augmented Dickey–Fuller (ADF) tests.

characterised as I(1) process with drifts. On the other hand, the results for the real interest rates series tend to favour the stationarity assumption, although the French R_{dt} seems to be better described by a trend stationary process.

3.2. Cointegration Analysis

There exists a large number of approaches to cointegration testing and estimation, ranging from Engle and Granger (1987) static regressions to various multivariate analyses which have a number of advantages in terms of the efficient use of the sample information and the underlying optimal inference that can be conducted. Among these, one of the most popular in empirical work is the Johansen (1991) MLE framework which assumes that the data are generated by a finite-order (linear) Gaussian VAR model. Although one may argue that the finite-order VAR model could stem from an approximate Wald representation of a non-linear process, it appears that, given our theoretical set-up, this assumption should be used with caution. A possible alternative is to use asymptotically median-unbiased estimators that do not require specific parametric representations of the short-run dynamics and that nevertheless lead to optimal inference. In this paper we use the fully modified ordinary least squares estimators (FMOLS) proposed by Phillips and Hansen (1990) and Hansen (1992a) based on semi-parametric corrections for endogeneity and serial correlation which in our case would stem from the presence of $\Delta \ln c_{mt}$ and $\Delta \ln c_{dt}$ in the right-hand side of equation (12).[9] This FMOLS estimator yields asymptotically optimal estimates of the non-stationary components[10] and is asymptotically equivalent to full information maximum likelihood parametric estimators for a rather large class of innovations processes. The conditions under which FMOLS estimators display these asymptotic optimal properties essentially reduce to the existence of some higher-order moments and of some mixing properties. Although it seems difficult to demonstrate formally that the process underlying the right-hand side of equation (12) satisfies these conditions required for the applicability of these non-parametric corrections, we will follow Ogaki and Park (1994) and given the form of the right-hand side of equation (12), we presume that these fully-modified approaches keep their optimal asymptotic properties in our case. As a by-product, usual asymptotic theory can be used to conduct inference on the cointegrating vector parameters. For example, the significance of a linear trend in the long-run relation implied by equation (12) can be tested using fully modified Wald or t-test statistics.

As is well-known in static cointegration regressions, any variable can theoretically be used as the regressand. Following Clarida (1994) and Ceglowski (1991) we decided to use $\ln c_{mt}$ as regressand. Table III reports the cointegration results for France and the USA. Standard errors are reported in parentheses. As pointed out by Haug (1996) and Cappucio and Lubian (1994), the way in which we estimate the long-run covariance matrix used to correct the estimates can play an important role in finite samples. We therefore report the FMOLS results for different choices of the kernel: QS, Parz. and Bart., denoting respectively Quadratic Spectral, Parzen and Bartlett kernels (see Andrews and Monahan, 1992). All our estimates are computed using an automatic plug-in bandwidth parameter after VAR prewhitening, which avoids the arbitrariness of choosing *a priori* the order of the truncation parameter. L_c is Hansen's (1992b) Lagrange

[9] Consequently, while super-consistent, we expect OLS static regressions to suffer from second-order asymptotic bias (see Phillips and Hansen, 1990) in the sense that the asymptotic distribution of the normalized bias is non-central.

[10] One should note, however, that this implies that the regressors of our problem form a set of full rank I(1) processes. If the latter assumption is violated, i.e. if the set also includes several I(0) variables, then, in the linear case at least, the Fully Modified GIVE or GMM estimators recently proposed by Kitamura and Phillips (1997) should be preferred.

Table III. Cointegration results — fully modified OLS results

$$\ln c_{mt} = \text{cst} + \frac{1}{\alpha_m}\ln\left(\frac{p_{dt}}{p_{mt}}\right) + \frac{\alpha_d}{\alpha_m}\ln c_{dt} - \frac{\mu_d - \mu_m}{\alpha_m}\text{trend} + \text{error}_t$$

Country	Kernel	cst	$-(\mu_d - \mu_m)/\alpha_m$	$1/\alpha_m$	α_d/α_m	L_c	$H(0, 1)$	*SupF*	*MeanF*
France	QS	−0·2373	0·0032	0·1560	3·3431	0·341	6·762*	7·759	3·585
		(0·060)	(0·001)	(0·141)	(0·267)				
	Parz.	−0·2380	0·0032	0·1592	3·3376	0·341	6·744*	7·587	3·573
		(0·061)	(0·001)	(0·142)	(0·268)				
	Bart.	−0·2380	0·0032	0·1530	3·3416	0·341	6·815*	7·794	3·603
		(0·060)	(0·001)	(0·141)	(0·267)				
	QS	−0·0824	—	0·4085	3·9689	0·290	—	6·866	2·890
		(0·007)		(0·134)	(0·113)				
	Parz.	−0·0824	—	0·4081	3·9683	0·297	—	6·982	2·949
		(0·007)		(0·133)	(0·113)				
	Bart.	−0·0824	—	0·4069	3·9707	0·291	—	7·013	2·908
		(0·007)		(0·133)	(0·112)				
USA	QS	−0·0404	−0·0003	0·7510	2·8523	0·275	0·022	12·884	5·359
		(0·121)	(0·002)	(0·062)	(0·442)				
	Parz.	−0·0436	−0·0002	0·7514	2·8396	0·277	0·015	12·866	5·419
		(0·120)	(0·002)	(0·061)	(0·438)				
	Bart.	−0·0456	−0·0002	0·7507	2·8336	0·290	0·011	12·817	5·430
		(0·121)	(0·002)	(0·062)	(0·441)				
	QS	−0·0582	—	0·7497	2·7876	0·103	—	8·533	1·818
		(0·008)		(0·061)	(0·053)				
	Parz.	−0·0582	—	0·7510	2·8523	0·104	—	8·381	1·786
		(0·008)		(0·060)	(0·052)				
	Bart.	−0·0582	—	0·7510	2·8523	0·103	—	8·462	1·798
		(0·007)	—	(0·060)	(0·053)				
5% crit values		With trend				0·778	3·84	17·30	7·69
		Without trend				0·690		14·80	6·17

Multiplier test which is based on the constancy of the intercept of the cointegration regression. This test can be interpreted as a test for the null of cointegration against the alternative of no-cointegration. $H(0, 1)$ is Park's (1990) Wald test for the null of deterministic cointegration computed on the residuals from the FMOLS regressions (see Park, 1990; Haug, 1996).[11] It has an asymptotic chi-square distribution with 1 degree of freedom under the null hypothesis of deterministic cointegration. Finally, the columns *SupF* and *MeanF* are statistics derived by Hansen (1992b) to test for the constancy of the parameters. Asymptotic critical values for L_c, *SupF*, and *MeanF* are taken from Hansen (1992b). We return to these statistics later in Section 4.

Table III shows some notable differences between the results obtained for France and for the USA which call for several comments. Notice, however, that contrary to the simulation outcomes reported in Haug (1996), the results remain almost unchanged if Parzen or Bartlett's kernels are used instead of a Quadratic spectral kernel.

[11] Park's canonical cointegration regressions were also computed but are not reported since they almost exactly coincide with the results obtained from Phillips–Hansen's FMOLS estimator.

Let us first consider the USA. As shown by the point estimates, standard errors as well as both the L_c and $H(0, 1)$ statistics, we cannot reject the null hypothesis of deterministic cointegration. This implies that the cointegration restriction derived in the preceding section seems to hold with the additional restriction that $\mu_d = \mu_m$ which stems from the insignificance of the trend term (i.e. the deterministic cointegration restriction). The last two columns show that one cannot reject the null hypothesis of parameter constancy of the long-run relation.

The results for France are less straightforward to interpret. When a linear trend is included in the cointegration regression, it appears from both the point estimate of $-(\hat{\mu}_d - \hat{\mu}_m)/\hat{\alpha}_m$ and from Park's (1990) $H(0, 1)$ that we may reject the deterministic cointegration restriction at any reasonable significance level while stochastic cointegration is not rejected[12] by L_c. Although in accordance with the result of Hall (1988) stating that there is no strong evidence in favour of a positive $1/\alpha$ for aggregate consumption in the USA, the resulting $1/\hat{\alpha}_m$ is surprisingly low and insignificant using a fully modified t-test. On the other hand, estimation of the model without trend, where we thus impose deterministic cointegration ($\mu_d = \mu_m$), provides more realistic point estimates. This is a typical situation where, although in both cases we cannot reject the null of stochastic cointegration, arguments based on economic theory tend to favour the results generated by a model which seems statistically misspecified. Consequently, we consider both cases in the second step of our analysis in order to investigate the sensitivity of the estimation of the habit-formation parameters to the maintained hypothesis about deterministic cointegration. Notice again that there is no sign of parameter non-constancy, irrespective of the assumption about the presence of a linear trend. We should observe that the approach we follow implicitly assumes that there is only one cointegrating vector. To check for this maintained hypothesis, we computed Johansen's Trace test for the number of cointegrating vectors. Although not reported here, the outcome of these tests do not provide any evidence in favour of a second cointegrating vector for both countries. Moreover, in both cases, the null hypothesis of weak exogeneity of $\ln c_{dt}$ and $\ln(p_{mt}/p_{dt})$ for the cointegrating vector parameters is rejected by the data, which is consistent with our theoretical model.

3.3. Short-run Analysis (Method)

We now proceed to the estimation of equation (6), in which the curvature parameters are restricted to their point estimates obtained from the cointegration analysis. The robustness of the cointegration results to the choice of the kernel and the Monte Carlo evidence reported in Andrews and Monahan (1992) and Cappucio and Lubian (1994) lead us to select the point estimates obtained with the QS kernel. Given the non-linear dynamic rational expectations formulation of the theoretical model, the non-linear IV version of GMM seems a natural method for estimating the remaining parameters of the Euler equations. In analogy to Engle and Granger (1987) two-step method, we assume that the asymptotic properties of the second-step GMM procedure are not affected by the first step estimation since the estimators for α_m and α_d from cointegrating regressions converge faster than the GMM estimators. The advantages of pursuing a cointegration analysis to identify and estimate the preference parameters from the utility function was first pointed out and discussed in detail by Ogaki (1992) and Ogaki and Park (1994). Basically, these advantages are all related to a substantial gain in robustness against several

[12] A similar rejection of non-cointegration is observed if one computes standard ADF or Phillips–Ouliaris tests using OLS cointegration regression.

potential problems such as measurement errors, short-run dynamic misspecification, preference shocks, which are known to affect the GMM (Ogaki, 1993a).

Defining $\theta_m = \theta\,\exp(\hat{\mu}_m)$ and $\theta_d = \theta\,\exp(\hat{\mu}_d)$, we use the point estimate of the trend coefficient, $\hat{\mu}_i$, to impose:

$$\theta_d = \theta_m\,\exp(\hat{\mu}_d - \hat{\mu}_m)$$

Denoting $\hat{\alpha}_i$ the point estimates obtained from the cointegration analysis, dividing equation (6) by $q_{it}^{-\hat{\alpha}_i}$ and using equation (1) leads to the following estimable form:

$$1 - \gamma_i\theta_i\left(\frac{c_{it} - \gamma_i c_{it-1}}{c_{it+1} - \gamma_i c_{it}}\right)^{\hat{\alpha}_i} - \frac{\theta_i}{R_{it+1}}\left(\left(\frac{c_{it} - \gamma_i c_{it-1}}{c_{it+1} - \gamma_i c_{it}}\right)^{\hat{\alpha}_i} - \gamma_i\theta_i\left(\frac{c_{it} - \gamma_i c_{it-1}}{c_{it+2} - \gamma_i c_{it+1}}\right)^{\hat{\alpha}_i}\right) = \varepsilon_{it+1} \tag{14}$$

for $i = m,\ d$ and where

$$\begin{aligned}\varepsilon_{it+1} = {} & -\gamma_i\theta_i\left(\frac{c_{it} - \gamma_i c_{it-1}}{c_{it+1} - \gamma_i c_{it}}\right)^{\hat{\alpha}_i} - \frac{\theta_i}{R_{it+1}}\left(\left(\frac{c_{it} - \gamma_i c_{it-1}}{c_{it+1} - \gamma_i c_{it}}\right)^{\hat{\alpha}_i} - \gamma_i\theta_i\left(\frac{c_{it} - \gamma_i c_{it-1}}{c_{it+2} - \gamma_i c_{it+1}}\right)^{\hat{\alpha}_i}\right) \\ & + \mathrm{E}_t\left[\gamma_i\theta_i\left(\frac{c_{it} - \gamma_i c_{it-1}}{c_{it+1} - \gamma_i c_{it}}\right)^{\hat{\alpha}_i} - \frac{\theta_i}{R_{it+1}}\left(\left(\frac{c_{it} - \gamma_i c_{it-1}}{c_{it+1} - \gamma_i c_{it}}\right)^{\hat{\alpha}_i} - \gamma_i\theta_i\left(\frac{c_{it} - \gamma_i c_{it-1}}{c_{it+2} - \gamma_i c_{it+1}}\right)^{\hat{\alpha}_i}\right)\right]\end{aligned}$$

which involves only stationary variables. When the two error terms are evaluated at the true value of the parameters, we have by assumption

$$\mathrm{E}[(\varepsilon_{mt+1}\varepsilon_{dt+1})' \,|\, \Omega_t] = 0$$

Let I_t be a vector including stationary variables taken from Ω_t that are observable by the econometrician. The moment restrictions used for the GMM estimation of the parameters can be summarized as

$$\mathrm{E}[(\varepsilon_{mt+1}\varepsilon_{dt+1})' \otimes I_t] = 0$$

The two equations are thus estimated jointly with the adequate cross-restrictions. Since the error terms include both c_{it+1} and c_{it+2}, they display a MA(1) structure and the instruments are to be lagged once more than in a standard lifecycle model. (Without habit formation, ε_{mt} and ε_{dt} are serially uncorrelated.)

As discussed in Hall (1993) and Ogaki (1993a), the GMM often appears to be sensitive to the chosen instrument set. In particular, for a fixed sample size, increasing the number of instruments increases the number of useful overidentifying restrictions but, on the other hand, may introduce substantial bias in the estimates of the coefficients. Accordingly, we define the following vectors of instruments: a minimalist one, containing a constant and the lagged interest rate, a medium-sized one, and a wider one including past levels of expenditures growth:

$$I_{1t} = (\text{constant}, r_{t-1})'$$

$$I_{2t} = \left(\frac{p_{dt-1}}{p_{dt-2}}, \frac{p_{mt-1}}{p_{mt-2}}, \text{trend}, \text{trend}^2, R_{mt-1}, R_{dt-1}\right)'$$

$$I_{3t} = \left(\frac{p_{dt-1}}{p_{dt-2}}, \frac{p_{mt-1}}{p_{mt-2}}, \text{trend}, \text{trend}^2, R_{mt-1}, R_{dt-1}, \frac{c_{mt-1}}{c_{mt-2}}, \frac{c_{dt-1}}{c_{dt-2}}\right)'$$

As suggested by Kocherlakota (1990) and Nelson and Startz (1990), we iterate on the weighting matrix (i.e. the inverse of the covariance matrix of the orthogonality conditions) in order to improve the properties of the estimators in our small sample. For France, two different estimates of the curvature parameters are used: with and without trend in the long run. Indeed, from the first step, we know that the trend is statistically significant. However, the economic interpretation of the trend may seem difficult. For this reason, we also estimate the model with the point estimates of α_m and α_d when we impose $\mu_d = \mu_m$ (no trend). Note that $\theta_m = \theta_d$ in the estimations without trend.

3.4. Short-run Analysis (Results)

Table IV presents the results. Standard errors are reported in parentheses. These are built on the basis of the heteroscedastic-consistent covariance matrix of Newey and West (1987). J is

Table IV. GMM estimates

Country	tr	I_t	n	θ_m	γ_m	γ_d	J_{test}	$LR_{\gamma m=\gamma d=0}$	$SupLR$
France	y	I_{1t}	1	0·959	0·58	0·41	1·23	4·84	N.A.
				(0·11)	(0·12)	(0·16)	[0·27]	[0·09]	
	y	I_{2t}	11	1·034	0·42	0·23	12·73	7·50	85·9
				(0·03)	(0·14)	(0·12)	[0·31]	[0·02]	
	y	I_{3t}	15	1·046	0·28	0·04	16·5	5·01	102·4
				(0·02)	(0·12)	(0·08)	[0·35]	[0·08]	
	n	I_{1t}	1	0·951	0·71	0·52	0·57	3·56	N.A.
				(0·08)	(0·14)	(0·20)	[0·45]	[0·17]	
	n	I_{2t}	11	0·999	0·52	0·27	12·5	11·4	85·1
				(0·01)	(0·16)	(0·14)	[0·33]	[0·00]	
	n	I_{3t}	15	1·001	0·38	0·04	17·2	8·73	108·4
				(0·01)	(0·12)	(0·08)	[0·31]	[0·01]	
USA	n	I_{1t}	1	0·958	0·42	0·74	0·23	1·89	N.A.
				(0·047)	(0·26)	(0·14)	[0·63]	[0·39]	
	n	I_{2t}	11	0·987	0·36	0·53	3·12	10·7	3·24
				(0·006)	(0·13)	(0·13)	[0·99]	[0·00]	
	n	I_{3t}	15	0·991	0·23	0·55	10·53	10·1	16·0
				(0·005)	(0·08)	(0·09)	[0·79]	[0·01]	
5% crit. values								5·99	13·2

Hansen's (1982) test for overidentifying restrictions, asymptotically χ^2 distributed with n degrees of freedom, where n is the number of overidentifying restrictions. Corresponding P-values are reported in brackets. $LR_{\gamma^m=\gamma^d=0}$ is a quasi-likelihood ratio test for the absence of habit formation, i.e. for H_0: $\gamma = \gamma_d = 0$. As suggested by Gallant (1987), it is computed as the normalized difference between the constrained objective function and the unconstrained one. The constrained estimation is computed with the weighting matrix provided by the unconstrained estimation. *SupLR* is the supremum of the sequence of the quasi-likelihood ratio type test for parameter constancy suggested by Andrews (1993). The critical values and the full sequence of these likelihood ratio test statistics are presented and discussed in the next section.

The main conclusions of the GMM estimation are the following:

- Hansen's J test measures the extent to which the residuals are effectively orthogonal to the instrument set. It can be seen as a global specification test. The numbers of degrees of freedom equals the number of restrictions imposed by the orthogonality conditions. These restrictions are not rejected at the 5% level for the two countries and the three different instrument sets.
- Concerning the estimates, $\gamma_i \varepsilon^{-g_i} \approx \gamma_i$ can be interpreted as the ratio of habit with respect to total consumption at steady state. On the basis of t-tests for these parameters γ_i, the habit-formation process appears significant in most cases and does not seem much affected by the first step (trend/notrend). Habit formation seems quantitatively more important for French imports and US domestic goods. The magnitude of their point estimates seems, however, to decline when we increase the number of instruments. Notice also that the point estimates of the γ_i could possibly be biased downwards in the presence of some durability effects in non-durables due, for example, to inventory holdings.
- The sensitivity of the estimation results to the choice of instruments set is well documented in the literature and seems especially evident for France for which γ_d even becomes insignificant when I_{3t} is used. Although we want to avoid making any definitive statement from this sensitivity, these results, coupled with the analysis of the constancy of parameters (see below), seems to point out that the retained specification may not be satisfactory for France.
- The significance of the habit-formation process based on the individual t-statistics is confirmed by the quasi-likelihood ratio test statistics. Indeed, the standard lifecycle model without habit formation is in most cases rejected at the 5% level.
- There is a tradeoff between (1) a low point estimates of φ_m (which includes both the discount factor and the parameter μ_d or μ_m so that it has not to be lower than 1) and a high level of habit formation and (2) high point estimates of θ_m and low habits.

Since we now have estimates for the γ_i's and the θ_i's it is possible to compute an approximation of the two IES following equation (8). These are reported and commented in the final section.

3.5. Analysis of the GMM Residuals

As seen from the specification of the utility function, we have limited the introduction of habit formation to a specification with a lag of one period. This is a crucial assumption of our model which may or may not limit its validity. Hence, it seems important to address this issue explicitly. A first possibility would be to modify equation (1) by introducing additional lagged c_{it-j} with $j > 1$ and then compute a quasi-LR test for the validity of a first-order habit formation. As long as we work with a finite number of additional lags, GMM could still, at least theoretically, be applied. As pointed out by Heaton (1995), these additional terms imply a larger MA structure in the error term of the Euler equations and in this case the estimation of the asymptotic covariance matrix of the estimator becomes difficult and the quality of the instruments (which have consequently to be lagged) may seriously deteriorate. A solution to this problem could be to rely on Simulated Methods of Moments as in Heaton (1995). This possibility lies, however, outside the scope of this paper and is left for further research.

One of the implications that can usefully be exploited to assess the validity of the specification of the retained simple habit-formation formulation is to recognize that under the validity of our theoretical specification, the bivariate process $(\varepsilon_{mt}, \varepsilon_{dt})'$ should behave like a first-order Vector Moving Average process, a VMA(1). Consequently, the implied univariate processes for $(\varepsilon_{mt}, \varepsilon_{dt})'$ should also be consistent with a MA(1) representation. One way to check for the potential misspecification of the short-run dynamic could therefore be to adopt the following strategy: (1) fit univariate MA(1) processes for both $\hat{\varepsilon}_{mt}$ and $\hat{\varepsilon}_{dt}$, (2) test for remaining dynamics in the fitted residuals of these MA(1) processes.

Table V reports some statistics on the residuals of these MA(1) models. The asterisk indicates that the corresponding statistic is significant at a 5% level. The column I_t indicates the instrument set used in the GMM estimation while 'tr' denotes whether or not a trend was included in the cointegration regression. $LB(q)$ is Ljung–Box statistic for qth-order serial correlation in the residuals of the MA(1) models computed with Diebold's (1986) correction for autoregressive conditional heteroscedasticity. $LB(q)$ is $\chi^2(q)$ distributed under the null of absence of serial correlation. $ARCH(q)$ is a standard Lagrange Multiplier test for qth-order ARCH effects in the residuals, $\chi^2(q)$ distributed under the null of absence of ARCH effects. *Norm* is Bera–Jarque's Normality test, $\chi^2(2)$ distributed under the null of Normality.

Graphical inspection shows that the assumption of homoscedastic innovations is hard to maintain. This is confirmed by the outcomes of standard Lagrange Multiplier tests for the absence of ARCH effects which point out the presence of conditional heteroscedasticity in almost

Table V. Misspecification tests

Country	tr	I_t		$LB(4)$	$LB(12)$	*Norm*	$ARCH(2)$	$ARCH(4)$
France	y	I_{1t}	$\hat{\varepsilon}_{mt}$	9·07	13·34	20·6*	0·13	13·6*
			$\hat{\varepsilon}_{dt}$	6·23	23·32*	15·7*	0·02	5·74
	y	I_{2t}	$\hat{\varepsilon}_{mt}$	5·02	12·51	23·1*	0·07	9·78*
			$\hat{\varepsilon}_{dt}$	7·81	18·63	4·2	0·00	0·58
	y	I_{3t}	$\hat{\varepsilon}_{mt}$	4·01	10·74	31·3*	0·11	6·10
			$\hat{\varepsilon}_{dt}$	5·94	13·80	2·7	0·39	0·65
	n	I_{1t}	$\hat{\varepsilon}_{mt}$	8·42	18·02	8·50*	0·08	10·96*
			$\hat{\varepsilon}_{dt}$	8·08	30·13*	10·2*	1·75	8·66
	n	I_{2t}	$\hat{\varepsilon}_{mt}$	5·93	14·42	5·9	0·14	6·66
			$\hat{\varepsilon}_{dt}$	9·25	22·49*	0·6	1·50	4·02
	n	I_{3t}	$\hat{\varepsilon}_{mt}$	4·15	11·30	8·0*	0·20	6·07
			$\hat{\varepsilon}_{dt}$	7·12	16·00	1·1	1·89	4·08
USA	n	I_{1t}	$\hat{\varepsilon}_{mt}$	2·26	11·23	1253·0*	10·98*	12·66*
			$\hat{\varepsilon}_{dt}$	9·39	27·66*	30·7*	8·95*	9·34
	n	I_{2t}	$\hat{\varepsilon}_{mt}$	4·40	12·54	1008·0*	8·94*	10·25*
			$\hat{\varepsilon}_{dt}$	9·48	25·02*	12·9*	8·13*	8·83
	n	I_{3t}	$\hat{\varepsilon}_{mt}$	4·68	12·99	560·0*	4·37	5·33
			$\hat{\varepsilon}_{dt}$	9·73*	26·33*	13·6*	8·87*	9·88*

all our residuals series, for the USA at least. Hence, we used Diebold's (1986) corrected Ljung–Box statistics in order to avoid spuriously significant serial correlation. As expected from the *LM* tests for ARCH effects, the assumption of Normality is rejected for all the US cases.

These test statistics point out that not much dynamics remain in the residuals of these MA(1) models. It is, for example, interesting to note that the few significant corrected Ljung–Box statistics are essentially those of the second equation, that is, those for $\hat{\varepsilon}_{dt}$, the innovation of the Euler describing the short-run dynamic behaviour of the domestic consumption. Nevertheless, although significant twelfth-order LB statistics are found (but not fourth-order statistics), the statistics are never very far from the 5% boundary and are mostly not significant if the chosen level was 1% (which is 26·23 for a $\chi^2(12)$ and 13·28 for a $\chi^2(4)$). One potential extension of this work (which is left for future research) is to investigate whether richer (non-linear) forms of habit formation, as, for example, in Campbell and Cochrane (1995) or Lettau and Uhlig (1995), could not improve or modify our conclusions.

Overall, in view of the moderate size of the serial correlation, we do not expect too serious an inconsistency to arise in our GMM estimation.

4. PARAMETER CONSTANCY ANALYSIS

Lucas (1976) argued that the parameters of traditional macroeconometric models depend crucially on agents' expectations and are unlikely to remain stable in a changing economic environment. In response to this critique, recent econometric practice has focused on the estimation of rational expectations models that have an explicit structural interpretation — Euler equations in particular. Thus, a natural criterion for judging the success of empirical Euler equations is the constancy of their 'deep', structural parameters. Given the two-step analyses retained in this paper, the investigation of the constancy of the structural parameters is also pursued in two different steps. Effectively, both the long-run and the short-run parameters should be constant over the retained sample period if our theoretical model is to be considered as a valid representation (explanation) of household behaviour. The potential non-constancy of the curvature parameters is investigated using appropriate formal statistical tests in cointegration regression. We follow the approach proposed by Hansen (1992b), based on FMOLS estimation, under the null hypothesis of the existence of a unique cointegration vector with constant parameters. Three different test statistics are considered under the assumption that the location of the potential break point is unknown. The first test statistic is in the spirit of traditional Chow tests: we compute a standard Chow F-statistic for a fixed break date t/T and then consider the sequence of statistics by varying the location of the break. The final statistics is then the supremum of this sequence.

$$SupF = \sup_{t/T \in [0{\cdot}15,\, 0{\cdot}85]} F_{t/T}$$

Under H_0, $SupF$ depends on both the number of variables in the cointegration regression and on the specification of the deterministic components. Asymptotic critical values from Hansen (1992b) are reported in Table III. From the sequence of $F_{t/T}$, Hansen (1992b) also proposes to compute the average value of the $F_{t/T}$. While the null hypothesis remains the same, the $MeanF$ is likely to be more powerful against gradual changes in the parameters. Finally, as shown in

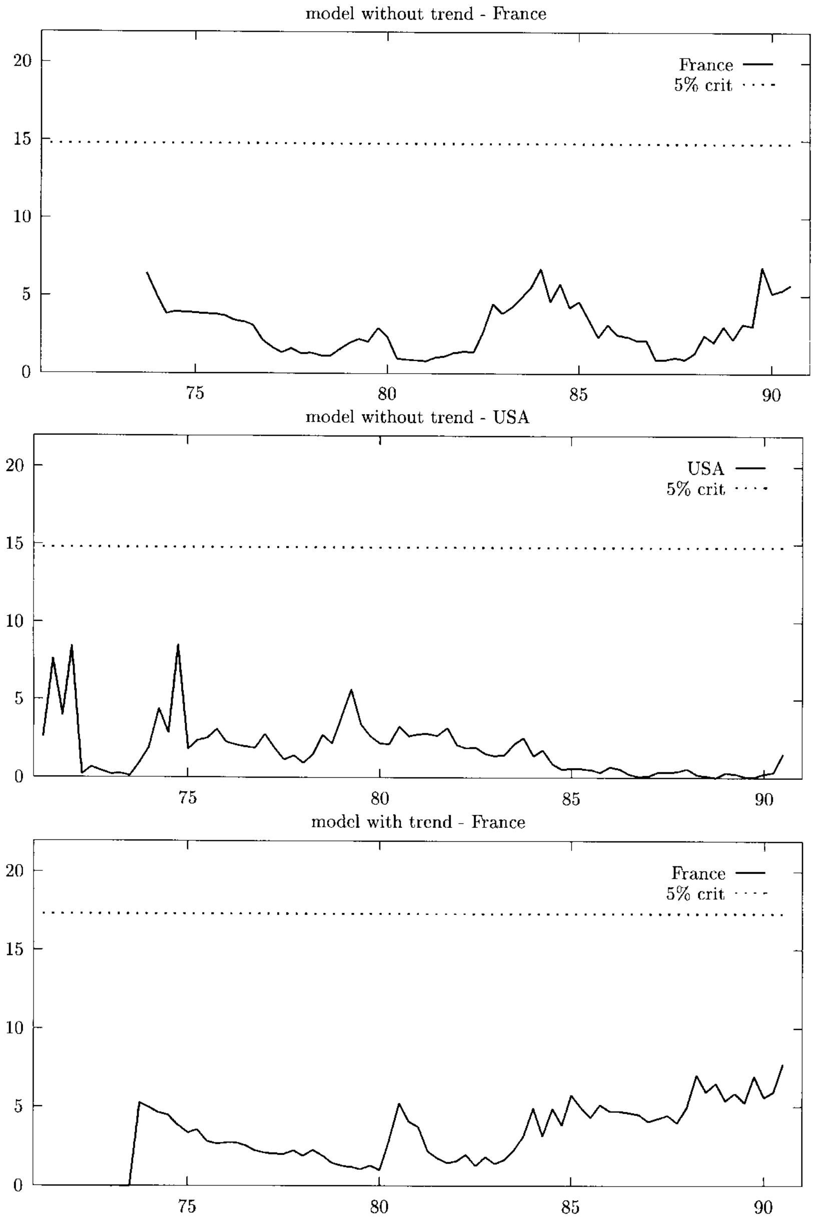

Figure 1. Constancy of the cointegration parameters

Table III, we also compute Hansen's L_c test statistic for parameter constancy against martingale variation in the constant term of the cointegration regression. Although formally built as a test for parameter constancy, L_c is easily interpreted as testing the null of cointegration. Figure 1 reports the sequence of $F_{t/T}$ over the interval [0·15, 0·85]. From these figures, we may not reject the constancy of the long-run parameters for both countries. This provides an additional argument in favour of the long-run implications of the theoretical model.

Given that we cannot reject the constancy of the long-run parameters, we may analyse the constancy of the short-run parameters conditionally on this. The analysis considers a sequence of *LR* tests (see Andrews, 1993), computed as the difference between the partial-sample GMM objective function evaluated at the full-sample GMM and at the partial sample GMM estimators. The structural break is allowed to occur in the interval[13] of time [0·25, 0·75]. The test can be performed for the two larger instrument sets which provide enough overidentifying restrictions. The critical value from Andrews (1993) is provided in Table IV and Figure 2 presents the sequences of the $LR_{t/T}$ statistics. We first note that the hypothesis of parameter constancy is rejected for France whatever instrument sets are used. For the USA, we observe some moderate[14] non-constancy with the larger set I_{3t}. However, it is likely that I_{3t} is too large given the size of the sub-sample used in the computation of $LR_{t/T}$. With I_{2t}, which is less subject to the above criticism, the parameter constancy cannot be rejected for the USA.

To evaluate the role of habit formation, we have also tested the parameter constancy of the standard lifecycle model (without habits) for the USA. This is reported in the third panel of Figure 2. In that case, the 5% critical value for the test is 7·93 instead of 13·16. Clearly the constancy of the parameters is now strongly rejected. The parameters seem to have experienced a shift during the period 78–81, which may correspond to the change in monetary policy of these years or to the second oil shock. This shows that the introduction of habit formation is important in obtaining a well-specified model of imports, at least for the USA.

An interesting implication of parameter constancy is the apparent robustness of the US results to some exogenous shocks that may have occurred during the retained sample period. In particular, although the oil price shocks are often believed to have had severe effects on consumption behaviour, they do not, for the USA at least, seem to induce significant changes in the parameters. Whether such shocks may explain the observed non-constancy in France is an interesting issue in itself but that may require further analysis. In particular, as shown by Hall and Sen (1996), the sequence of statistics we use for the detection of parameter non-constancies may, in case of rejection of the null, reflect either a change in the parameters and/or a change in the functional form. Disentangling these two effects is, however, outside the scope of this paper. Although not reported, the partial-sample estimates of the parameters for France indicate that the rate of time preference increases substantially at the end of the period. Among the potential explanations of this phenomenon, we could argue that our model should display endogenous discount rates. Indeed, a rise in the discount rate is consistent with the idea of Uzawa (1968) that a higher level of consumption implies a higher rate of time preference. This is, however, difficult to sustain *a priori*, since we usually think that the rich are more patient than the poor. An alternative explanation could be a too high aggregation level. It should be interesting to investigate whether different goods have experienced very different evolutions in their prices, and whether the weight of these goods in the consumer basket has changed over the sample period. Effectively, observed parameter non-constancy may reveal heterogeneity in preferences, in goods or in initial wealth endowments (Fève and Langot, 1995).

[13] For numerical reasons linked to the non-linear structure of our second-step problem, it was necessary to reduce the interval of time compared to the one used in the first step.
[14] The constancy is rejected at 5%, but not at 1%.

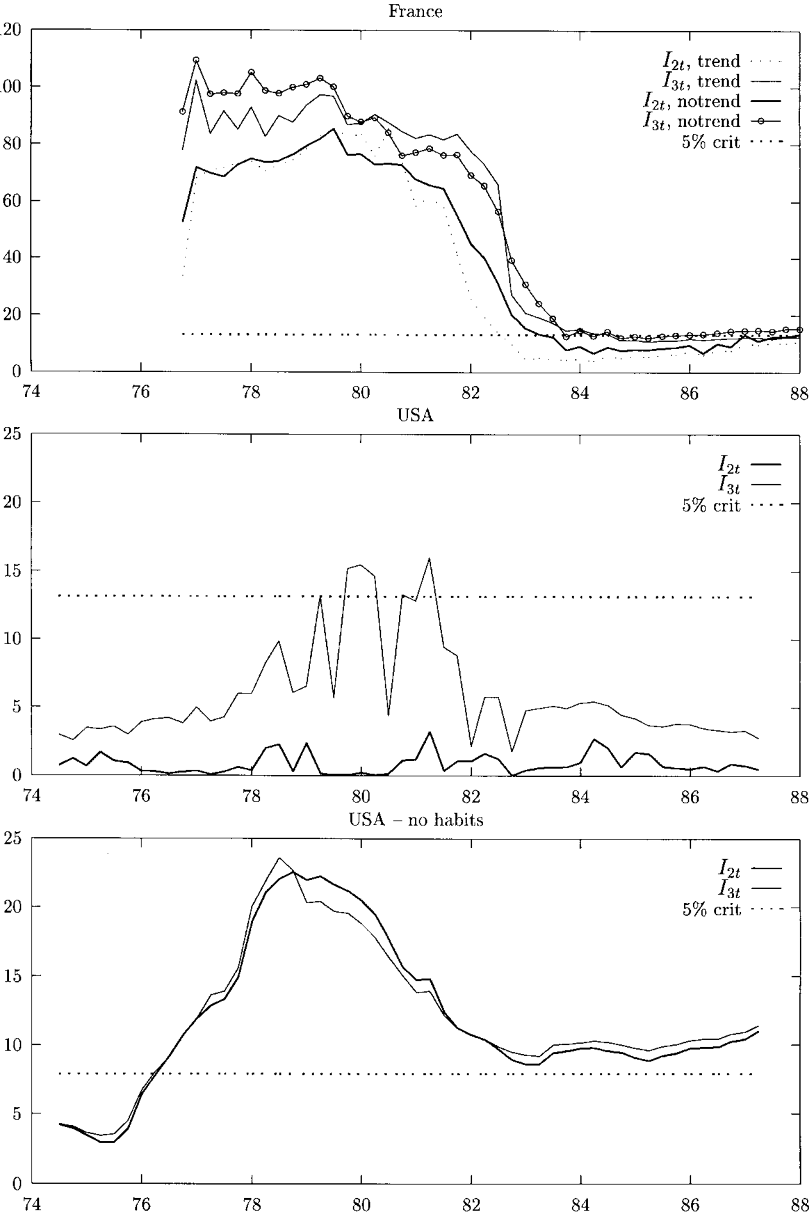

Figure 2. Constancy of the parameters estimated by GMM

5. CONCLUSIONS

The purpose of this paper was to study non-durable imports demand, by extending previous work by Clarida (1994) and Ceglowski (1991). We considered a two-good version of the lifecycle model in which we introduce time-non-separability in households' preferences. The model is estimated using quarterly time-series data for the USA and France. Using the information in the

Table VI. Comparisons

	Homogeneous goods			Imports and domestic goods			
	Hall USA 47–83	Fers.–Const. USA 47–89	Ogaki–Park USA 48–86	Ceglowski USA 67–88	Clarida USA 67–90	This paper USA 67–94	This paper France 70–94
$1/\hat{\alpha}$	0·10	0·67	0·43				
$1/\hat{\alpha}_m$				0·89	0·95	0·75	0·41
$1/\hat{\alpha}_d$				0·33	0·44	0·27	0·10
IES	0·10	0·32	0·27				
IES_m				0·89	0·95	0·27	0·07
IES_d				0·33	0·44	0·05	0·05
$\hat{\gamma}$		0·28	0·19				
$\hat{\gamma}_m$						0·36	0·53
$\hat{\gamma}_d$						0·53	0·27

$\hat{\alpha}$ and $\hat{\gamma}$ are parameters from a model with only one consumption good.
IES is the approximation of the intertemporal elasticity of substitution using equation (8).

observed stochastic and deterministic trends, we derive a cointegration restriction used to estimate curvature parameters of the instantaneous utility function. The remaining parameters are estimated in a second step by GMM. Table VI compares our results (without trend in the long run) with those obtained in other studies, all using the same class of utility functions, with or without habits and with one or two goods.

Let us first consider the parameters $1/\alpha_i$, which are related to the willingness of consumers to shift consumption across time in response to changes in interest rates. The bulk of empirical evidence suggests that this parameter lies around or below unity. An exception is Hall (1988) who concludes that the IES is unlikely to be much above 0·1 for non-durable aggregate consumption goods (in the USA). However, since the paper by Constantinides (1990) it is known that habit formation introduces a gap between $1/\alpha$ and the IES. He evaluates that the IES is of the order of one fourth the value of $1/\alpha$. In their study of food consumption, Naik and Moore (1996) evaluate this gap on the order of one half. When habit formation is allowed, Ferson and Constantinides (1991) and Ogaki and Park (1994) show that a relatively low IES can be made compatible with a relatively high value of $1/\alpha$. Comparing our results with those existing for import demand, our estimates confirm a large and significant $1/\alpha_m$ for the USA, three times larger than $1/\alpha_d$. Following Ceglowski (1991), one implication of this result is the importance of real interest rates in determining the consumer demand for imports. As pointed out in the Introduction, this contrasts sharply with standard import equations. This conclusion should be moderated by the argument that $1/\alpha_i$ is not a good measure of intertemporal substitution when habit formation is significant. If we use the corrected measure proposed by Boldrin, Christiano, and Fisher (1995) (see equation (8)), the short-run intertemporal elasticity of substitution in the USA is equal to $1/\hat{\alpha}_i$ divided by 3 for imported goods and almost by 6 for domestic goods. This implies that the gap between the two goods is still larger but, of course, the absolute magnitude of the short-run intertemporal effect is reduced by habit formation.

For France, the parameters $1/\alpha_i$ are less important, but the one for imports is still four times larger than that for domestic goods. Thus, the French estimates confirm that imports are more sensitive to changes in interest rate than domestic goods. This holds even after the correction computed from equation (8).

Although the finding of a relatively large elasticity of intertemporal substitution of non-durable imports as compared to domestic consumption cannot easily be accounted for by consumer theory, one potential explanation could be the lower share of necessities in non-durable imports.[15] An interesting empirical finding is the constancy of the curvature parameters for both countries.

Considering now the implications of the model in terms of non-linear dynamics, an important result is that the overidentifying restrictions implied by the habit-formation assumption are not rejected by the data for the two countries. For France, the estimation of the habit-formation process is, however, not fully convincing. (The point estimates of the parameters are sensitive to the chosen instrument set. In some cases one finds a relatively weak effect of habits. The constancy of the short-run parameters is rejected.) For the USA, the habit-formation process seems particularly significant. On the one hand, the model without habit formation is strongly rejected on the basis of quasi-likelihood ratio tests. On the other, the introduction of habits is useful for the constancy of the short-run parameters. These results support the view that ignoring habits may help in explaining the frequent rejection of the lifecycle hypothesis.

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REFERENCES

Amano, R. and T. Wirjanto (1996), 'Intertemporal substitution, imports and the permanent income model', *Journal of International Economics*, **40**, 439–457.

Andrews, D. (1993), 'Test for parameter instability and structural change with unknown change point', *Econometrica*, **61**, 821–856.

Andrews, D. and J. McDermott (1995), 'Nonlinear econometric models with deterministically trending variables', *Review of Economic Studies*, **62**, 343–360.

Andrews, D. and C. Monahan (1992), 'An improved heteroscedasticity and autocorrelation consistent covariance matrix', *Econometrica*, **60**, 953–966.

Argyle, M. (1987), *The Psychology of Happiness*, Methuen, London.

Asseery, A. and D. Peel (1991), 'Estimates of a traditional aggregate import demand model for five countries', *Economic Letters*, **35**, 435–439.

Boldrin, M., L. Christiano and J. Fisher (1995), 'Asset pricing lessons for modeling business cycles', Working Paper, NBER.

Campbell, J. and J. Cochrane (1995), 'By force of habits: a consumption-based explanation of aggregate stock market behavior', Working Paper, NBER.

Canova, F., M. Finn and A. Pagan (1994), 'Evaluation of a real business cycle model', in C. Hargreaves (ed.), *Nonstationary Time Series Analysis and Cointegration*, Oxford University Press, Oxford.

Cappucio, N. and D. Lubian (1994), 'Fully modified estimation of cointegrating vectors via var pre-withening: a simulation study', Working Paper, University of Padua.

[15] The question whether the inventory behaviour of the importers can be responsible for the higher intertemporal elasticity of substitution of imports is left for future research.

Ceglowski, J. (1991), 'Intertemporal substitution in import demand', *Journal of International Money and Finance*, **10**, 118–130.

Clarida, R. (1994), 'Cointegration, aggregate consumption, and the demand for imports: A structural econometric investigation', *American Economic Review*, **84**, 298–308.

Constantinides, G. (1990), 'Habit formation: a resolution of the equity premium puzzle', *Journal of Political Economy*, **98**, 519–543.

Dickey, D. and W. Fuller (1979), 'Distribution of the estimators for autoregressive time series with a unit root', *Journal of the American Statistical Association*, **74**, 427–431.

Diebold, F. (1986), 'Testing for serial correlation in the presence of arch', Working Paper, University of Pennsylvania.

Duesenberry, J. (1949), *Income, Saving, and the Theory of Consumer Behavior*. Harvard University Press, Cambridge, MA.

Eichenbaum, M., L. Hansen and K. Singleton (1988), 'A time series analysis of representative agent models of consumption and leisure choice under uncertainty', *Quarterly Journal of Economics*, **103**, 51–78.

Engle, R. and C. Granger (1987), 'Co-integration and error correction: representation, estimation and testing', *Econometrica*, **55**, 251–276.

Ferson, W. and G. Constantinides (1991), 'Habit persistence and durability in aggregate consumption', *Journal of Financial Economics*, **29**, 199–240.

Fève, P. and F. Langot (1995), 'La méthode des moments généralisées et ses extensions', *Economie et Prévision*, **119**, 139–169.

Gagnon, J. (1988), 'Adjustments costs and international trade dynamics', *Journal of International Economics*, **26**, 327–344.

Gallant, D. (1987), *Nonlinear Statistical Models*, Wiley, New York.

Goldstein, M. and M. Khan (1985), 'Income and price elasticities in foreign trade', in R. Jones and P. Kenen (eds), *Handbook of International Economics, Vol. 2*, North-Holland, Amsterdam.

Hall, A. (1993), 'Some aspects of generalized method of moments estimation', in G. Maddala, C. Rao and H. Vinod (eds), *Handbook of Statistics, Vol. 11*, Elsevier, New York.

Hall, A. and A. Sen (1996), 'Structural stability testing in models estimated by generalized method of moments', Working Paper, North Carolina State University.

Hall, R. (1988), 'Intertemporal substitution in consumption', *Journal of Political Economy*, **96**, 339–357.

Hansen, B. (1992a), 'Tests for parameter instability in regressions with I(1) processes', *Journal of Business and Economic Statistics*, **10**, 321–335.

Hansen, B. (1992b), 'Efficient estimation and testing for cointegration vector with the presence of deterministic trends', *Journal of Econometrics*, **53**, 81–121.

Hansen, L. (1982), 'Large sample properties of generalized method of moments estimators', *Econometrica*, **50**, 1029–1054.

Hansen, L. and K. Singleton (1982), 'Generalized instrumental variables estimation of nonlinear rational expectations models', *Econometrica*, **50**, 1269–1286.

Haug, A. (1996), 'Tests for cointegration: a Monte Carlo comparison', *Journal of Econometrics*, **71**, 89–115.

Heaton, J. (1995), 'An empirical investigation of asset pricing with temporally dependent preference specifications', *Econometrica*, **63**, 681–717.

Helson, H. (1964), *Adaptation-level Theory: An experimental and systemic approach to behavior*, Harper and Row, New York.

Husted, S. and T. Kollintzas (1987), 'Linear rational expectations equilibrium laws of motion for selected U.S. raw material imports', *International Economic Review*, **28**, 651–670.

Johansen, S. (1991), 'Estimation and hypothesis testing of cointegration in vector Gaussian autoregressive models', *Econometrica*, **59**, 1551–1580.

Kim, Y. (1993), 'Frisch demand functions and intertemporal substitution in consumption', *Journal of Money, Credit and Banking*, **25**, 445–454.

Kitamura, Y. and P. Phillips (1997), 'Fully modified iv, give and gmm estimation with possibly non-stationary regressors and instruments', *Journal of Econometrics*, **80**, 85–123.

Kocherlakota, N. (1990), 'On tests of representative consumer asset pricing models', *Journal of Monetary Economics*, **26**, 285–304.

Kollintzas, T. and S. Husted (1984), 'Distributed lags and intermediate goods imports', *Journal of Economic Dynamic and Control*, **8**, 303–327.

Lettau, M. and H. Uhlig (1995), 'Can habit formation be reconciled with business cycle facts?' Working Paper, CentER (Tilburg).

Lucas, R. (1976), 'Econometric policy evaluation: a critique', in K. Brunner and A. Metzler (eds), *The Phillips Curve and Labour Market*, Carnegie-Rochester Conference Series on Public Policy, North-Holland, Amsterdam.

MacKinnon, J. (1991), 'Critical values for cointegration', in R. Engle and C. Granger (eds), *Long-run Economic Relationships: Readings in Cointegration*, Oxford University Press, New York.

Mah, J. (1994), 'Japanese import demand behaviour: the cointegration approach', *Journal of Policy Modeling*, **16**, 291–298.

McLaughlin, K. (1995), 'Intertemporal substitution and λ — constant comparative static', *Journal of Monetary Economics*, **35**, 193–213.

Muellbauer, J. (1988), 'Habits, rationality and myopia in the life cycle consumption function', *Annales d'Economie et de Statistique*, **9**, 47–70.

Naik, N. and M. Moore (1996), 'Habit formation and intertemporal substitution in individual food consumption', *The Review of Economics and Statistics*, **78**, 321–328.

Nelson, C. and R. Startz (1990), 'The distribution of the instrumental variable estimator and its *t*-ratio when the instrument is a poor one', *Journal of Business*, **63**, 125–140.

Newey, W. and K. West (1987), 'A simple positive definite, heteroskedasticity and autocorrelation consistent covariance matrix', *Econometrica*, **55**, 703–708.

Ogaki, M. (1992), 'Engel's law and cointegration', *Journal of Political Economy*, **100**, 1027–1046.

Ogaki, M. (1993a), 'Generalized method of moments: Econometric applications', in G. Maddala, C. Rao and H. Vinod (eds), *Handbook of Statistics, Vol. 11*, Elsevier, New York.

Ogaki, M. (1993b), 'Unit roots in macroeconomics: A survey', *BOJ Monetary and Economic Studies*, **2**, 131–154.

Ogaki, M. and J. Park (1994), 'A cointegration approach to estimating preference parameters', *Journal of Econometrics*, forthcoming.

Otto, G. (1992), 'Testing a present-value model of the current account: Evidence from US and Canadian time series', *Journal of International Money and Finance*, **11**, 414–430.

Park, J. (1990), 'Testing for unit roots and cointegration by variable addition', *Advances in Econometrics*, **8**, 107–133.

Phillips, P. and B. Hansen (1990), 'Statistical inference in instrumental variables with I(1) processes', *Review of Economic Studies*, **57**, 99–124.

Phillips, P. and P. Perron (1988), 'Testing for a unit root in time series regression', *Biometrika*, **75**, 335–346.

Schmidt, P. and P. Phillips (1992), 'LM tests for a unit root in the presence of deterministic trends', *Oxford Bulletin of Economics and Statistics*, **54**, 257–287.

Shapiro, M. (1986), 'Dynamic demand for capital and labour', *Quarterly Journal of Economics*, **101**, 513–542.

Urbain, J.-P. (1992), 'Error correction models for aggregate imports', in M. Dagenais and P.-A. Muet (eds), *International Trade Modelling*, Chapman and Hall, London.

Urbain, J.-P. (1995), 'Partial versus full system modelling of cointegrated systems: An empirical illustration', *Journal of Econometrics*, **69**, 177–210.

Uzawa, H. (1968), 'Time preference, the consumption function and optimum asset holdings', in J. Wolfe (ed.), *Value, Capital and Growth: Papers in Honour of Sir John Hicks*, University of Edinburgh Press.

Winder, C. and F. Palm (1996), 'Stochastic implications of the life cycle consumption model under rational habit formation', *Recherches Economiques de Louvain*, **62**, 403–412.

