Back to the Future? Habits and Rational Addiction in UK Tobacco and Alcohol Demand

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Abstract

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JEL Classification Codes: I12, D12, C32

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Abstract

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I Introduction

Rational addiction as defined by Becker and Murphy (1988) implies that consumption of addictive goods with negative health implications is still consistent with forward-looking maximization of utility from stable preferences. Addiction is rational in the sense that consumers go beyond the pure myopic behavior (habit persistence and reinforcement from past consumption) and anticipate the consequences of future consumption. This theory has allowed economists to treat demand for addictive goods, previously disregarded as irrational, and has been tested fairly successfully on alcohol and tobacco consumption. Empirical tests of the theory involve the estimation of demand models allowing for a response of demand to past levels of consumption and current and future prices (Becker, Grossman and Murphy, 1994; Chaloupka, 1991; Grossman and

Chaloupka, 1998; Baltagi and Griffin, 2001). While some authors (Suranovic et al., 1999) have objected that such forward-looking behavior ignores potential adverse health consequences in the future, the persistence of unhealthy behavior has been justified with adjustment and withdrawal costs that prevent consumers switching to a healthier consumption bundle (Jones, 1999).

This paper develops a dynamic modeling approach for the Almost Ideal Demand System (AIDS), where the forward-looking hypothesis is combined with that of (convex) adjustment costs in the presence of non-stationary cointegrated variables. While the task of validating the theory for a single equation demand model has been accomplished in the above cited literature, to our knowledge there are no studies at a system level. Our application investigates aggregate consumption of alcoholic beverages and tobacco in the UK over the 1963-2003 period, using quarterly data. Accounting for interaction between alcohol and tobacco demand through a systemwise approach is not a trivial extension, since it is acknowledged that the cross-price effect in alcohol and tobacco demand might be quite relevant and complementarity between the two goods is a probable outcome (Decker and Schwartz, 2000).

The rational habit hypothesis is explicitly introduced within an AIDS modeling framework in Alessie and Kapteyn (1991), Andrikopoulos et al. (1997), Weissemberg (1986) and Rossi (1987). These two latter studies take the forward-looking perspective of habit formation into explicit account. Our dynamic demand system is derived from an intertemporal optimization problem involving quadratic adjustment costs, where the AIDS model is used to represent the equilibrium targets that consumers would pursue in the absence of adjustment costs. The resulting forwar-looking model reads as an error-correcting model.

As observed in Johnson et al. (1992), the error-correcting nature of demand systems is perceived as a key feature when variables are non-stationary. Cointegrated demand system have often proved consistent with the underlying theoretical restrictions, otherwise subject to frequent rejection (see Keuzenkamp and Barten, 1995). In our set-up the resulting error-correcting demand system is consistent with both myopic and forward-looking behavior.

In some respects, our formulation is similar to Weissemberger's (1986) intertemporal model. However, some differences are introduced. First, we allow for second-order adjustment costs, which means that costs of adjustment are experienced by modifying both the expenditure levels and their rates of change over time, in order to achieve the optimal expenditure allocation. This should allow for a more flexible dynamic adjustment structure, as compared to specifications based only on first-order adjustment costs, the so-called "polynomial frictions" (Kozicki and Tinsley, 1999). Second, we embed a classic version of the AIDS within the dynamic model, which is assumed to represent equilibrium demand in the absence of frictions. Third, we account for non-stationarity and cointegration in the variables and propose a simple two-stage procedure for estimating the dynamic demand system. In the first stage cointegration techniques are applied to obtain consistent estimates the parameters of the static (long-run) AIDS relationship. Having fixed the parameters of the AIDS at their super-consistent first-step estimates, the Generalized Method of Moments (GMM) is implemented in the second stage to estimate the remaining parameters, which relate to the dynamic adjustment structure represented by a system of interrelated Euler equations. Any existing econometric package can be used to implement the proposed method.

The paper is organized as follows. In Section II we introduce the model through a two-step specification, the long run and the adjustment dynamics. In Section III we discuss the estimation procedure and in Section IV the propsed demand model is applied to quarterly UK alcohol and tobacco data. Section V draws some conclusions.

II The model

We consider a representative consumer who undertakes a two-stage decision process. In the first stage he decides the optimal (target) expenditure allocation across the different goods. In the second stage this target behavior is embedded into a quadratic cost of adjustment-disequilibrium framework.

A Long-run equilibrium

The target level of consumption is assumed to follow the linear approximation of the static and flexible AIDS model of Deaton and Muellbauer (1980). This implies that the long-run equilibrium is free from habits and expectations and reflects the Marshallian demand function derived from an utility-maximizing consumer with a price-independent generalized logarithmic (PIGLOG) expenditure function. The equilibrium relationship is given by the following system of demand equations:

$$w_{it}^{*} = \gamma_{i0} + \sum_{j=1}^{n} \gamma_{ij} \log P_{jt} + \lambda_i \log\left(\frac{Y_t}{P_t}\right) + \delta_i t \tag{1}$$

$$i = 1, ..., n \quad , \quad t = 1, ...T$$

where n is the number of goods, w_{it}^{*} is the aggregate target level for the i-th expenditure share as predicted by consumer theory,

$$w_{it} = (P_{it}Q_{it} / \sum_{j=1}^{n} P_{jt}Q_{jt}) = (P_{it}Q_{it}/Y_t)$$

is the actual expenditure share for good i at time t, P_{jt} is the price of good j at time t, Q_{it} is the quantity of good i purchased at time t, Y_t is the total expenditure at time t and the non-linear price index P_t can be adequately approximated by the Stone index $\log P_t^{*} = \sum_{h=1}^{n} w_{kt} \log P_{ht}$. A linear trend with coefficient δ_i was included in the model to capture systematic trends (e.g. smooth structural changes) in demand patterns.

System (1) is linear in the preference parameters γ_{i0}, γ_{ij} and λ_i and is obtained under the assumption of intertemporal separability. In order to respect the underlying theoretical assumptions, the following restrictions must hold in (1):

$$\sum_{i=1}^{n} \gamma_{i0} = 1, \quad \sum_{i=1}^{n} \gamma_{ij} = 0, \quad \sum_{i=1}^{n} \lambda_i = 0, \quad \sum_{i=1}^{n} \delta_i = 0 \tag{2}$$

$$\sum_{j=1}^{n} \gamma_{ij} = 0 \tag{3}$$

$$\gamma_{ij} = \gamma_{ji}. \tag{4}$$

These constraints represent respectively the adding up, homogeneity and symmetry assumptions from microeconomic theory. Given (2) the demand system is singular by construction. In order to avoid the related econometric problems, the usual procedure consists in dropping one equation from the system[1].

Following Ng (1995) and using some algebra the model can be parameterized as

$$w_{it}^* = \sum_{j=1}^{n-1} \gamma_{ij} \log\left(\frac{P_{jt}}{P_{nt}}\right) + \gamma_i^h \log(P_{nt}) + \lambda_i \left(\frac{Y_t}{P_t}\right) + \gamma_{i0} + \delta_i t \tag{5}$$

where $\gamma_i^h = \sum_{j=1}^{n} \gamma_{ij}$ and equilibrium expenditure shares are expressed in terms of relative prices and real total expenditure. The advantage of this formulation is that the homogeneity constraint (3) here corresponds to the restriction $\gamma_i^h = 0$. For the rest of the discussion we rewrite compactly (5) as:

$$w_t^* = \Gamma z_t + \gamma + \delta t \tag{6}$$

where $w_t^* = (w_{1t}^*, ..., w_{mt}^*)'$, $m = n - 1$, $z_t = (p_{1t}, ..., p_{mt}, p_{m+1t}, y_t)'$, $p_{jt} = \log(P_{jt}/P_{nt})$, $j = 1, ..., m$, $p_{nt} = \log(P_{nt})$, $y_t = \ln(Y_t/k_t P_t)$, $\Gamma = [\gamma_{ij} \vdots \gamma_i^n \vdots \lambda_i]$, $i = 1, ..., m$, $j = 1, ..., m-1$ is a $m \times (m+2)$ matrix

[1] As known, maximum likelihood estimates are invariant to the choice of the equation to be dropped (see Barten, 1969).

and $\gamma = (\gamma_{10}, ..., \gamma_{m,0})'$, $\delta = (\delta_1, ..., \delta_m)'$ are $m \times 1$ vectors.

B Dynamics

Consumers are unable to achieve equilibrium in each time period. This is due to habit persistence and the costs of adjusting the consumption bundle to meet future expectations. Let $\widetilde{w}_{t+j} = w^*_{t+j} + e_{t+j}$, where w^*_{t+j} is defined as in (6) and represents the equilibrium expenditure shares that minimize consumers utility costs in the absence of market frictions, and e_{t+j} is an $m \times 1$ error term added on the right-hand-side of the static AIDS to capture discrepancies between the agent's information set and the econometrician's one[2].

In each time period, the representative consumer solves the following cost-minimization problem

$$\min_{\{w_{t+j}\}} E_t \sum_{j=0}^{\infty} \rho^j [(w_{t+j} - \widetilde{w}_{t+j})' D_0 (w_{t+j} - \widetilde{w}_{t+j}) \\ + \Delta w_{t+j}{}' D_1 \Delta w_{t+j} + \Delta^2 w_{t+j}{}' D_2 \Delta^2 w_{t+j}] \quad (7)$$

where w_t, w_{t-1} and w_{t-2} are given at time t. In (7) $E_t \cdot = E(\cdot \mid \Omega_t)$ is the expectation operator conditional on the information set available at time t, $\Omega_{t-1} \subseteq \Omega_t$, $\Delta w_t = w_t - w_{t-1}$, $\Delta^2 w_t = \Delta w_t - \Delta w_{t-1}$, ρ $(0 < \rho < 1)$ is a time-invariant discount factor and D_0, D_1 and D_2 are $m \times m$ symmetric positive definite matrices. It is assumed that $\{w_t, w_{t-1}, ... \ z_t, z_{t-1}, ... \ e_t, e_{t-1} ...\} \subseteq \Omega_t$ and that $E_t e_{t+j} = 0$ for $j \geq 1$.

The first addendum in (7) measures the cost of not attaining the

[2] In other words the error e_t is assumed to be known to the representative consumer but unknown to the econometrician at time t, see e.g. Hansen and Sargent (1991).

long run target w^*_{t+j}, i.e. disequilibrium costs and the second and third addenda measure respectively the costs of changing w_t and Δw_t. In the rational addiction context, the disutility of adverse health consequences from consumption is accounted for in the first addendum, while the costs of adjustment (withdrawal) discussed by Jones (1999) are modeled in the remaining two terms.

If D_0, D_1 and D_2 are specified as non-diagonal matrices, cross-adjustment and cross-disequilibria costs arise. It is worth noting that differently from Weissemberger (1986), it is assumed that adjustment costs in (7) originate from quarterly rather than yearly changes in consumption allocations. This choice is motivated by the nature of consumption decisions we investigate in Section 4. It is reasonable to assume, indeed, that alcohol and tobacco consumers process new information more than only once a year[3].

The first-order necessary conditions to solve (7) are given by the system of (second-order) interrelated Euler equations

$$\rho^2 D_2 E_t \Delta^2 w_{t+2} - \rho[D_1 + 2D_2] E_t \Delta w_{t+1} + [D_1 + 2\rho D_2]\Delta w_t$$

$$+ D_2 \Delta^2 w_t + D_0(w_t - \widetilde{w}_t) = 0_{m\times 1} \quad (8)$$

and a set of transversality conditions[4]. Using (6) and simple algebra the Euler equations can be written as the expectations-based error-

[3]Alcohol and tobacco consumers probably process new information on a weekly basis. Our analysis is nevertheless limited by the available data frequency.

[4]The solution to the problem (7) is discussed in Pesaran (1991) for the case of a single decision variable and in Binder and Pesaran (1995) and Kozicky and Tinsley (1999) for the case of a multiple decision variable. Here we maintain that the conditions ensuring the existence of a unique stable solution to the rational expectations model (9) are fulfilled.

correcting model

$$E_t\Delta w_{t+2} = \rho^{-1}\Psi_1 E_t\Delta w_{t+1} - \rho^{-2}\Psi_2\Delta w_t$$

$$- \rho^{-2}\Delta w_{t-1} - \rho^{-2}\Upsilon(w_t - \Gamma z_t - \gamma - \delta t) + \varphi_t \quad (9)$$

where $\Psi_1 = [\Psi + (2+\rho)I_m]$, $\Psi_2 = [\Psi + 2\rho I_m]$, $\Psi = D_2^{-1}D_1$, $\Upsilon = D_2^{-1}D_0$ and $\varphi_t = \rho^{-2}\Upsilon e_t$. The matrices Ψ_1, Ψ_2 and Υ in (9) need not to be symmetric and the term $(w_t - w_t^*) = (w_t - \Gamma z_t - \gamma - \delta t)$ is the vector of expenditure share disequilibria, i.e. deviations of actual expenditure shares from the optimal long run levels that would prevail in the absence of frictions.

System (9) can be re-written as

$$\Delta w_t = \rho\Psi_2^{-1}\Psi_1 E_t\Delta w_{t+1} - \rho^2\Psi_2^{-1}E_t\Delta w_{t+2}$$

$$- \Psi_2^{-1}\Upsilon(w_t - \Gamma z_t - \gamma - \delta t) + \Psi_2^{-1}\Delta w_{t-1} + \varphi_t^* \quad (10)$$

where $\varphi_t^* = \rho^2\Psi_2^{-1}\varphi_t$. The formulation (10) of the system of Euler equations shows that the expenditure shares at time t depend on: (i) expected changes in expenditure shares one and two periods (quarters) ahead (forward-looking behavior); (ii) deviations of actual expenditure shares from equilibrium levels (disequilibria); (iii) changes in lagged expenditure shares (myopic habit persistence).

As Ψ_1, Ψ_2 and Υ are generally not diagonal, the pattern of expenditure shares for the i-th good depends on its own dynamics as well as on the dynamics of all goods in the system. In particular, the Υ matrix contains the own and cross-adjustment coefficients, i.e. the parameters in each row measure how expenditure shares react to the own disequilibrium as well as to the disequilibria involving the other goods. For simplicity, considering the case $m = 2$ (a three goods demand system $n = 3$ with $m = n - 1 = 2$ two modeled expenditure shares); then the Υ matrix is given by

$$\Upsilon = \begin{bmatrix} \omega_{11} & \omega_{12} \\ \omega_{21} & \omega_{22} \end{bmatrix}$$

where e.g. the structural parameter ω_{11} measures how Δw_{1t}, react to the own two periods lagged disequilibrium $(w_{1t-2} - w^*_{1t-2})$, whereas ω_{12} indicates whether Δw_{1t} react to the two periods lagged disequilibrium $(w_{2t-2} - w^*_{2t-2})$ characterizing the other expenditure share.

By solving the model (9) forward, it can be shown that the system is consistent with a dynamic specification where consumers react to lagged disequilibria and future expected changes of prices and total expenditure[5]. To show this, one can simply focus on the case where the D_2 matrix in (7) is zero, i.e. when consumers face first-order adjustment costs only. When $D_2 = 0_{m\times m}$ model (9) collapses to the system of (first-order) interrelated Euler equations

$$\Delta w_t = \rho E_t \Delta w_{t+1} - \Pi(w_t - \Gamma z_t - \gamma - \delta t) + \xi_t \tag{11}$$

where $\Pi = D_1^{-1} D_0$ and $\xi_t = \Pi e_t$. This model, which is nested in (10), can be solved forward as in e.g. Nickell (1984) and opportunely manipulated

[5]For a formal derivation see e.g. Kozicki and Tinsley (1999), Section 3 and Appendix A1.

to yield the optimal error-correcting decision rule

$$\Delta w_t = (I_m - \Lambda)(w_{t-1} - \Gamma z_{t-1} - \widetilde{\gamma} - \delta t) + \sum_{j=0}^{\infty} (\rho\Lambda)^j (I_m - \Lambda)\Gamma E_t \Delta z_{t+j} + \zeta_t \quad (12)$$

where Λ is a $m \times m$ matrix whose elements are opportunely related to that of Π in (11), $\zeta_t = (I_m - \rho\Lambda)(I_m - \Lambda)e_t$ and $\widetilde{\gamma} = \gamma + \delta$.[6] It is evident that changes in expenditure shares in (12) depend on the lagged disequilibria and expectations on future expenditure shares, prices and expenditure levels. This model represents a process of "rational" habit formation as the consumer depicted by (12) is forward as well as backward-looking.

Two further points about the specification outlined in this section should be made. First, model (9) embodies both the hypotheses of forward-looking behavior and convex adjustment costs, therefore it results in a tight dynamic structure.[7] Second, Heien and Durham (1991) argue that the habit effects modeled through lagged dependent variables are likely to be overstated when aggregate time-series data are used, as information about consumers heterogeneity is ignored. This results in higher residual autocorrelation, ultimately and incorrectly captured by the habit effect. This is also part of the usual aggregation argument, as

[6] As (11) and (12) are different representations of the same RE model, it can be shown that making inference on the adjustment parameters of (11) is equivalent to make inference on the parameters of (12) after having solved for the unknown future expectations on Δz_t variables and viceversa, provided that the relationship connecting the two parameterization is taken into explicit account, see e.g. Fanelli (2002).

[7] Quadratic costs can be restrictive, albeit mathematically convenient, therefore rejection of the model should not be intended as a clear cut evidence against the rational habit formation hypothesis as rational habit formation might still hold under a different dynamic structure.

individual behaviors are not necessarily reflected in aggregate models. Although this is certainly true, as most of the empirical research and policy analysis is based on aggregated time series data, it is certainly desirable to conduct a specification search to improve habit formation and rational addiction modeling.

III Estimation procedure

The unrestricted parameters of the system of Euler equations (9) are contained in (Γ, γ, δ, ρ, Ψ, Υ), where in particular (Γ, γ, δ) refers to the long-run AIDS and (ρ, Ψ, Υ) refers to the dynamic adjustment structure implied by the system of interrelated Euler equations. We propose a simple two-step procedure for estimating model (9) in presence of non-stationary cointegrated variables.

First step

We assume that w_t and z_t are I(1) and cointegrated such that the term $d_t = (w_t - \Gamma z_t - \gamma - \delta t)$ is I(0). Under these assumptions, the efficient estimation of (Γ, γ, δ) can be carried out through cointegration methods. Tests of the hypotheses of homogeneity and symmetry involve the elements in Γ and are characterized by chi-squared distributions[8].

The recent literature on the cointegrated AIDS model can be split into two main research streams relying on different estimation methods. The first bulk of works stems from Ng (1995) and Attfield (1997)[9], who adopt a triangular vector error correction (TVEC) representation

[8] Phillips (1991) and Johansen (1991) discuss the theoretical conditions under which efficient estimators that belong to the locally asymptotically mixed normal (LAMN) class can be obtained.

[9] See also Duffy (2003a) for an application to alcohol demand.

of the demand system and implement Dynamic Ordinary Least Squares (DOLS), see e.g. Stock and Watson (1993). The alternative modelling procedure is the Johansen (1996) full information maximum likelihood (FIML) technique, based on a Vector Error Correction (VEC) representation of the demand system as in e.g. Ben Kaabia and Gil (2001). When comparing empirically the two approaches, Fanelli and Mazzocchi (2002) show that the VEC has the advantage of providing a "natural" framework for testing (and framing) some of the hypotheses underlying the cointegrated AIDS, such as the exogeneity of prices and total expenditure[10].

Second step: GMM

The estimation of system (9) can be accomplished through GMM, provided that the matrix Γ, which contains the AIDS preference parameters, is replaced with the super (order T) consistent estimates obtained in the first step. Using the decomposition $\Delta w_{t+2} = E_t \Delta w_{t+2} + \eta_{t+2}$, where the rational expectations forecast error η_t is such that $E_t \eta_{t+2} = 0$ and $E_t \eta_{t+1} = 0$, by lagging the model by two time periods, the system of Euler equations (9) can be rewritten as

$$\Delta w_t = \rho^{-1} \Psi_1 \Delta w_{t-1} - \rho^{-2} \Psi_2 \Delta w_{t-2}$$

[10]Since the seminal paper of Summers (1959), it has been frequently argued that expenditure should not be assumed as exogenous in consumer demand systems without testing it. Attfield (1985) argued that the erroneous assumption of expenditure exogeneity could be another factor leading to the improper rejection of demand theory restrictions. Within aggregate demand systems it appears also reasonable to allow the presence of feedbacks from demand to prices, i.e. endogeneity of prices.

$$- \rho^{-2}\Delta w_{t-3} - \rho^{-2}\Upsilon\ \widehat{d}_{t-2} + u_t \qquad (13)$$

where $u_t = \eta_t - \rho^{-1}\Psi_1\eta_{t-1} + \varphi_{t-2}$ and $\widehat{d_t} = (w_t - \widehat{\Gamma}z_t - \widehat{\gamma} - \widehat{\delta}t)$ is the estimated disequilibria term obtained in the first-step. The matrices Ψ_1 and Ψ_2 are restricted as in (9).

Under the above hypotheses on the order of integration of variables the model (13) involves I(0) variables and reads as an error-correcting system non-linear in the parameters. The substitution of the "true" disequilibria $d_t = (w_t - \Gamma z_t - \gamma - \delta t)$ with the estimated $\widehat{d_t}$ does not affect the asymptotics (order $\sqrt{T}$) of the estimators of (ρ, Ψ, Υ) due to the super-consistency result.

It is assumed that the rational expectations forecast error η_t and the disturbance e_t are homoscedastic and $E(\eta_t e_t') \neq 0_{m\times m}$, implying that $E(\eta_t \varphi_t') \neq 0_{m\times m}$. In this case, the disturbance term u_t in (13) follows a MA(2)-type process and

$$E(u_t \mid \Omega_{t-3}) = 0_{m\times 1}. \qquad (14)$$

This orthogonality conditions in (14) are the basis for estimating (13) through GMM. Let $\theta = (\rho, vec(\Psi)', vec(\Upsilon)')'$ be the $a \times 1$ ($a = 1 + 2m^2$) vector containing the unrestricted parameters of (13) and θ_0 the "true" value of θ. From (14) it turns out that for any (stationary) vector s_t of dimension $q \times 1$ such that $s_t \in \Omega_{t-3}$ (i.e. dated $t-3$ and earlier) and $mq \geq a$,

$$E\left[h_t(\theta_0)\right] = 0_{mq\times 1} \qquad (15)$$

where $h_t(\theta_0) = u_t(\theta_0) \otimes s_t$ and $u_t(\theta)$ denotes the vector of disturbances of (13)[11]. The orthogonality conditions (15) can be employed to form a GMM estimator of θ by choosing $\widehat{\theta}_T$ as the solution to

$$\min_{\theta} \; h_T(\theta)' \; W_T \; h_T(\theta)$$

where $h_T(\theta) = \frac{1}{T}\sum_{t=1}^{T} h_t(\theta)$ is the $mq \times 1$ vector of sample moments and W_T is the weighting matrix. An optimal GMM estimator can be achieved by exploiting an heteroscedasticity and autocorrelation consistent (HAC) estimator of the covariance matrix as in Andrews (1991) or Newey and West (1994)[12]. The over-identifying restrictions test statistic, $J_T = T \; h_T(\widehat{\theta})' \; \widehat{W_T} \; h_T(\widehat{\theta})$, is χ^2 distributed with $mq - a$ degrees of freedom.

The system of Euler equations (13) is linear in parameters aside from the discount factor ρ. However, as it is generally difficult to estimate the discount factor within the class of forward-looking models we consider here, we adopt the common practice of pre-specifying it[13].

IV Results

The empirical analysis is based on a time series of T=161 quarterly observations of UK consumer expenditure on alcoholic beverages and tobacco over the period 1963:1 to 2003:1. All data are freely supplied on-line by the UK Office for National Statistics (www.statistics.gov.uk).

[11] It is clear that one could also consider different instruments in each of the m equations of the dynamic demand system.

[12] The estimation of model (13) through GMM leads to conventional questions about instrument choice, see e.g. Stock et al. (2002).

[13] Clearly plausible values of ρ are close to (but less than) one. Most studies find that variations in ρ do not significantly affect estimates of the other parameters.

Data on aggregate UK household consumer expenditure on alcoholic beverages are based on volume of sales and average prices of individual types of alcoholic beverages for "off-licence" trades. This information is obtained from a continuous survey of retail outlets. Estimates for tobacco are based on data obtained from Her Majesty's Customs & Excise (HMCE) relating to the quantities of tobacco released for sale within the UK. Quarterly household aggregated expenditure is obtained from several independent sources, including the Retail Sales Inquiry and the Expenditure and Food Survey (which merges the previous Family Expenditure Survey and the National Food Survey). Prices series are the Retail Price Index (RPI) for alcoholic beverages, tobacco and the all-item RPI. Prices and expenditure time series were scaled to be 1 at the mean point to reduce the bias from the Stone Index approximation.

The system includes two expenditure share equations, respectively for alcoholic beverages (w_{1t}) and tobacco (w_{2t}), while a numeraire equation (w_{3t}) for all remaining goods was dropped from estimation to overcome the singularity problem ($\sum_{i=1}^{3} w_{it} = 1$); hence $m = n - 1 = 3 - 1 = 2$. The information required to estimate the system is completed by the prices of alcoholic beverages, tobacco and all other goods p_{1t}, p_{2t} and p_{3t} and total expenditure y_t, after deflation through the Stone index. Prices and total expenditure are collected in the vector $z_t = (p_{1t}, p_{2t}, p_{3t}, y_t)'$.

Figure 1 shows the overall trend in alcohol and tobacco consumption (real expenditure at 1995 prices) over the sample period and the expenditure shares w_{1t} and w_{2t}. UK household consumption of alcohol has risen at a fairly constant rate, even if the expenditure share does not show the same trend due to the aggregate expenditure increase. The fall in tobacco consumption, especially from the 1980s, is evident from both the real expenditure and expenditure share graphs.

The first-step of our procedure starts with the estimation of the cointegrated AIDS (6). We used a VEC for $X_t = (w_{1t},\ w_{2t},\ z_t')'$ and the Johansen method (Johansen, 1996). The VEC includes a liner trend restricted to belong to the cointegration space and three centered seasonal dummies to capture seasonal patterns characterizing expenditure shares. The lag length was fixed at 5 as suggested by standard order selection criteria (AIC, SC, HQ) and residuals diagnostic tests. Computations were performed through PcGive10.0 and E-Views 4.0.

Before computing the Johansen Trace test for cointegration rank, we examined the roots of the characteristic polynomial associated with the VEC. The eigenvalues of the companion matrix associated with the VEC suggest the presence of unit roots at the long run frequency (as expected)[14]. The Trace test is reported in Table 1 and indicates the presence of four unit roots in the six-dimensional system X_t, corresponding to two cointegrating relations. Given the cointegration rank $r = m = 2$, FIML estimates of the preference parameters (Γ, γ, δ), are reported in Table 2.

The symmetry and homogeneity constraints characterizing the elements of Γ were tested through a likelihood ratio (LR) statistic and

[14] The eigenvalues of the companion matrix associated with the VAR suggest also the possibility of unit roots at the seasonal frequency (Hylleberg et al., 1990). A throughout analysis of the seasonal pattern of UK tobacco and alcohol demand goes well-beyond the purposes of the present paper. However, it is recognized in the specialized literature that to the extent that the dynamics of the VEC are correctly specified, the possible presence of unit roots at the seasonal frequencies does not pose any additional issue on usual cointegration tests at the zero frequency and the estimation of long run relationships.

not rejected. This is an encouraging result, in contrast with many applied studies, which could reflect a specification improvement. The weak and strong exogeneity of prices and total expenditure with respect to the structural parameters of the AIDS were sharply rejected[15]. The long-run relationships also show a negative and significant trend for both products, a shift of preferences away from the two addictive goods possibly due to increased information and health concerns.

Table 3 reports the estimates of the long-run Marshallian elasticities for the homogeneity and symmetry constrained system. Some results are striking, albeit plausible and consistent with previous studies. Demand for alcohol in the UK is quite price-elastic (-1.23) and there is clear evidence of complementarity between alcohol and tobacco consumption, as in the study by Jones (1989) for the UK and Decker and Schwartz (2000) on US individual data. In the long-run, there is also a very high expenditure elasticity, which contrasts the negative trend observed in preferences, as increasing incomes lead to higher consumption. Duffy (2003b) also finds some complementarity between tobacco and spirits and a high expenditure elasticity for spirits. However, his higher level of disaggregation allows to distinguish across alcoholic beverages, and beer and wine are found to be substitutes of tobacco. For tobacco consumption, our results differ from Duffy (2003b), as the Marshallian long-run own-price and expenditure elasticities are non-significant, showing a con-

[15] The "endogeneity" of total expenditure can be easily justified if one realizes that y_t represents by construction the denominator of expenditure shares. Observe that the violation of the weak exogeneity of z_t implies that the efficient estimation of the AIDS through Dynamic OLS (DOLS) should be performed by including a number of leads of Δz_t in addition to lags in the regression of w_t on z_t, see e.g. Ng (1995) for a throughout discussion.

sumer whose long-run consumption equilibrium is independent from the traditional economic factors. Habits and addiction are the main determinants of changes in consumption levels.

After replacing the preference parameters of the AIDS (Γ, γ, δ) with the estimates of Table 1 we next moved to the second-step estimation of the system of interrelated Euler equations (13). Details of the econometric specification are as follows. For GMM estimation, the set of stationary instruments was $s_t = (\Delta_4 w_{1t-3},\ \Delta_4 w_{2t-3},\ \Delta p_{1t-3},\ \Delta p_{2t-3},\ \Delta p_{3t-3},\ \Delta w_{1t-4},\ \Delta w_{2t-4},\ \Delta p_{1t-4},\ \Delta p_{2t-4},\ \Delta p_{3t-4},\ \Delta y_{t-4},\ \Delta w_{2t-5})'$ where for a given (logged) variable, v_t, $\Delta_4 v_t = v_t - v_{t-4} = (1-L^4)v_t$ generates yearly changes[16]. It can be easily recognized that the vector of instruments s_t belongs to the information set a time $t-3$ and earlier. The weight matrix was estimated through a HAC procedure with Bartlett weights and Newey and West's (1994) criterion for bandwidth. The discount factor was prefixed at 0.98 consistently with a quarterly average real discount rate of 2% (i.e. a real yearly rate of about 8%).

GMM estimates are summarized in Table 4 along with the J_T statistic for over-identifying restrictions. In Table 5 we summarized Shea's (1997) measures of instrument relevance[17]. The partial correlation among the

[16]The filter $(1-L^4)$ removes all unit roots characterizing v_t at all frequencies. Three centered seasonal dummies where included on the right hand side of (13) (and in the instrument list) to account for deterministic seasonal patterns characterizing the variables of the demand system. For simplicity estimates of coefficeints associated with seasonal dummies are not reported in Table 3.

[17]In models with one explanatory variable, the R^2 obtained from regressing the endogenous (explanatory) variables on the instrument vector can be considered a useful measure of instrument relevance. In multivariate models, however, one cannot measure relevance by simply regressing each explanatory variable on the instrument vector in turn; indeed, if instruments are highly collinear the R^2 might result high for each explanatory variable even when instruments are actually "weak". Shea's (1997) simple method allows to compute, for each explanatory variable, partial R^2 measures of instrument relevance, $\overline{R}_p^2$, by correcting opportunely for the correlation among

righ-hand side variables of system (13) and the selected list of instruments range from 0.30 to 0.53, which is a reasonable outcome for variables in their first differences (growth rates).

The estimates of Table 4 can be summarized in two important results. First, the estimated interrelated dynamic structure of adjustment suggests both backward and forward-looking behavior. Second, besides the complementarity of alcohol and tobacco discussed above, a further relevant link emerges in the consumption of these two addictive goods. Alcohol and tobacco adjust not only to their own past disequilibria, but also to each other's disequilibria as suggested by the estimated Υ matrix. This adjustment is faster (but less significant) for alcohol with respect to disequilibria in the tobacco relationship.

V Conclusions

Our paper suggests a dynamic specification for the Almost Ideal Demand System which is consistent with both backward and forward-looking behavior, with quadratic costs of adjustment. It is argued that this specification is consistent with the rational addiction hypothesis. Estimation is based on a two-step strategy, where cointegration techniques are used for estimating the long run demand system that would prevail in the absence of frictions and GMM for estimating the interrelated system of adjustment towards equilibria.

The empirical application, based on a 40-years long time series of UK alcohol, tobacco and other goods expenditure supports the specifi-

instruments. Observe that in Table 5 we did not report partial $\overline{R}_p^2$ for deterministic seasonal dummies.

cation choice and leads to four relevant results: (i) the joint backward and forward-looking specification is consistent with the proposed application; (ii) the theoretical restrictions of homogeneity and symmetry, often rejected in time series-based AIDS models, are also valid within the adopted specification; (iii) exogeneity of prices and total expenditure in the long-run relationship is strongly rejected and the proposed method allows to account for endogeneity, overcoming a relevant limit of most empirical applications; (iv) as few empirical studies have investigated before, there is a strong complementarity between alcohol and tobacco consumption behavior, not only in terms of price reactions, but also in adjusting to past cross-disequilibria.

These results, together with estimates of elasticities that show some noticeable differences from previous studies on similar data, have important policy implications, as models ignoring the extent of these dynamic, complementarity and endogeneity issues are unlikely to be adequate for simulations.

Our study also raises three relevant theoretical and empirical issues which still need to be addressed; (i) checking for the impact on results of alternative cost-adjustment structures; (ii) taking into account alternative estimation methods; and (iii), it may be difficult to generalize the results of an aggregate time-series study, as addictive behavior are known to vary significantly across socio-demographic segments and a panel data extension (as in Baltagi and Griffin, 2001) is desirable. The encouraging results of the proposed application suggest that dynamic issues deserve a careful, even if complex, treatment when addictive goods are considered and further study addressing the above limitations is desirable.

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FIGURES AND TABLES

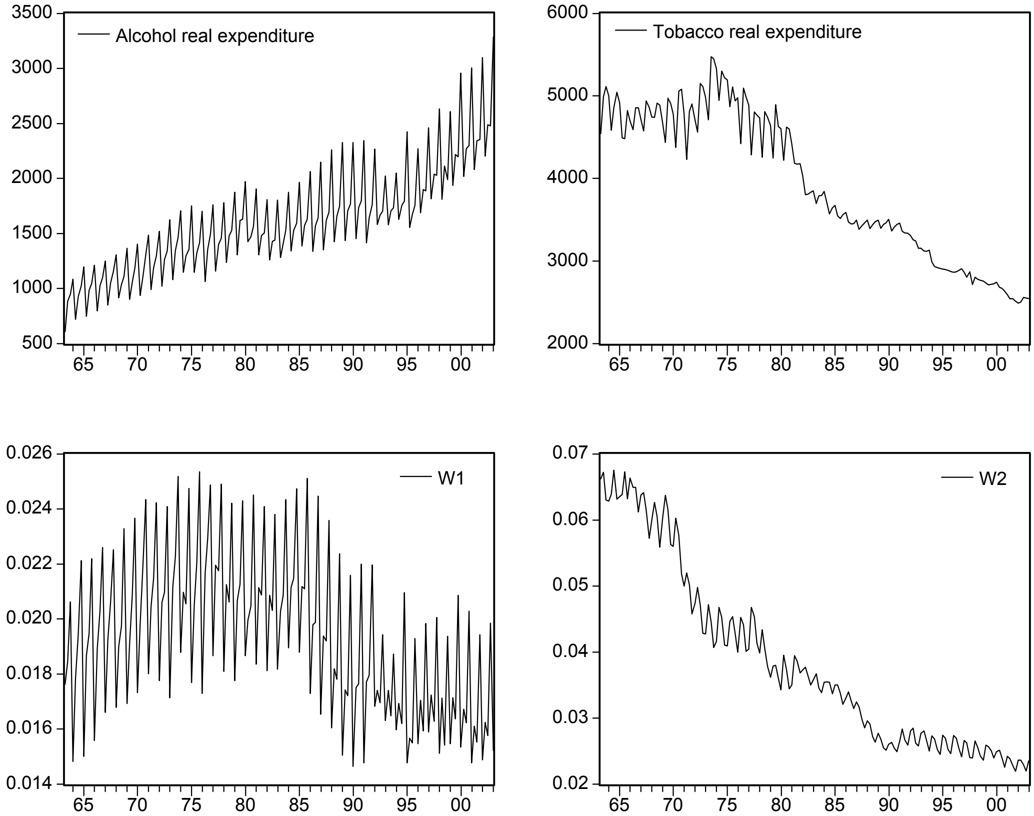

Figure 1: Trends in alcohol and tobacco consumption

Number of unit roots	Rank (r)	Trace	5% c.v.
6	0	125.2	114.9
5	1	90.9	89.9
4	2	59.7	62.6
3	3	32.7	42.2
2	4	13.4	24.4
1	5	4.7	12.4

Table 1: Johansen Trace test for cointegration rank based on the VEC with 5 lags, three centered seasonal dummies and a linear trend restricted to lie in the cointegration space. Notes: 5% asymptotic critical values are taken from Table 15.4 in Johansen (1996).

Commodity	p_{1t}	p_{2t}	p_{3t}	y_t	tr
w_{1t}^* (Alcohol)	-0.003 (-0.25)	-0.020 (-5.0)	0.024	0.054 (3.9)	-0.0003 (33.3)
w_{2t}^* (Tobacco)	-0.020 (-5.0)	0.039 (16.2)	-0.019	-0.039 (-6.5)	-0.0002 (-50.0)
LR: homogeneity and symmetry	$\chi^2(3)$=2.45 [0.48]				
LR: weak exogeneity of z_t	$\chi^2(8)$=40.20 [0.00]				
LR: strong exogeneity of z_t	$\chi^2(40)$=82.77[0.00]				

Table 2: Upper panel: FIML estimates of the parameters of the AIDS (7) with homogeneity and symmetry imposed. Lower panel: Likelihood ratio (LR) tests for the hypotheses of homogeneity and symmetry and on the weak and strong exogeneity of prices and total expenditure. Notes: the cointegration rank is fixed at 2 consistently with the results in Table 1; t-statistics in round brackets, p-values in squared brackets.

		Elasticities		
Commodity	Alcohol	Tobacco	Other Goods	Expenditure
w^*_{1t} (Alcohol)	-1.23 (0.42)	-1.16 (0.05)	1.21	3.80 (0.56)
w^*_{2t} (Tobacco)	-0.50 (0.01)	0.01 (0.01)	-0.48	-0.01 (0.02)

Table 3: Marshallian Long-Run Elasticities. Standard errors between brackets

GMM estimates of system (13)

$$\rho = \underset{\text{fixed}}{0.98}\ , \quad \widehat{\Psi}_1 = \begin{bmatrix} \underset{(10.34)}{1.48} & \underset{(5.38)}{1.03} \\ \underset{(-10.83)}{-1.62} & \underset{(-15.13)}{-0.88} \end{bmatrix}\ , \quad \widehat{\Psi}_2 = \begin{bmatrix} \underset{(10.34)}{0.46} & \underset{(5.38)}{1.03} \\ \underset{(-10.83)}{-1.62} & \underset{(-15.13)}{-1.90} \end{bmatrix}$$

$$\widehat{\Upsilon} = \begin{bmatrix} \underset{(-2.96)}{-0.004} & \underset{(-1.50)}{-0.117} \\ \underset{(-2.49)}{-0.005} & \underset{(-2.95)}{-0.244} \end{bmatrix}\ , \quad J_T = \underset{[0.70]}{12.57}$$

Alcohol equation:	Tobacco equation:
$R^2 = 0.93$	$R^2 = 0.63$
s.e. =0.005	s.e. = 0.006

Table 4: GMM estimates of the parameters of the system of Euler equationsl (13) with J_T test for over-identifying restrictions. Notes: the first row inf the matrices $\widehat{\Psi}_1$, $\widehat{\Psi}_2$ and $\widehat{\Upsilon}$ refers to the alcohol equation whereas the second row refers to the tobacco equation; the vector of instruments is $s_t = (\Delta_4 w_{1t-3}, \Delta_4 w_{2t-3}, \Delta p_{1t-3}, \Delta p_{2t-3}, \Delta p_{3t-3}, \Delta w_{1t-4}, \Delta w_{2t-4}, \Delta p_{1t-4}, \Delta p_{2t-4}, \Delta p_{3t-4}, \Delta y_{t-4}, \Delta w_{2t-5})'$ plus three deterministic seasonal dummies; t-statistics in round brackets below parameters estimates; the p-value associated withr the J_T test is in squared bracket and is computed from a χ^2 distribution with $mq - a$ =16 degree of freedom, with m =2, q =15 the number of instruments in s_t (including three deterministic seasonal dummies), and a = 14 is the number of free estimated parameters (including those associated with the three deterministic seasonal dummies).

Regressors of system (13)	$\overline{R}_p^2$
Δw_{1t-1}	0.48
Δw_{2t-1}	0.53
Δw_{1t-2}	0.31
Δw_{2t-2}	0.49
$\widehat{d}_{1t-2}$	0.49
$\widehat{d}_{2t-2}$	0.30

Table 5: Shea's (1997) $\overline{R}_p^2$ partial measures of instrument relevance, see footnote 17 for details. Notes: the vector of instruments is $s_t = (\Delta_4 w_{1t-3},\ \Delta_4 w_{2t-3},\ \Delta p_{1t-3},\ \Delta p_{2t-3},\ \Delta p_{3t-3},\ \Delta w_{1t-4},\ \Delta w_{2t-4},\ \Delta p_{1t-4},\ \Delta p_{2t-4},\ \Delta p_{3t-4},\ \Delta y_{t-4},\ \Delta w_{2t-5})'$ plus three deterministic seasonal dummies.