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ECONOMETRIC MODELING OF CONSUMERS' EXPENDITURE IN VENEZUELA

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ABSTRACT

Starting from a theoretical model with optimizing economic agents, we develop a highly parsimonious econometric model of consumers' expenditure on non-durables and services in Venezuela for 1970-85. Disposable income, liquidity, and inflation determine expenditure in an economically sensible fashion. The empirical model is robust and has constant, well-determined parameter estimates. In specifying it, econometric methodology plays a fundamental role, and we address issues of empirical model design and evaluation, cointegration, exogeneity, policy analysis, and encompassing. Using the last concept, a large class of expectations and VAR models is found to be incompatible with the data. In particular, Hall's (1978) hypothesis (derived from the life cycle-permanent income hypothesis) that expenditure is a random walk and only predictable from its own past is firmly rejected. The empirical model provides a clear interpretation for why that is so.

Econometric Modeling of Consumers' Expenditure in Venezuela

Julia Campos and Neil R. Ericsson[1]

1. Introduction

Modeling consumers' expenditure has been a major focus of theoretical and empirical economic research over the last several decades. In spite of those efforts, considerable gaps remain between theoretical and empirical models and, among the empirical models, major disagreements exist as to the determinants of consumers' expenditure. One possible cause for those gaps and disagreements is the <u>econometric</u> methodology adopted. Typically, an economic theory is developed, economic series corresponding to the theoretical constructs are found, the theoretical model is estimated with those data, and various economic hypotheses are tested on the empirical model. While theory, choice of data, and estimation are important parts of empirical modeling, they are not sufficient to ensure that inferences based on the empirical model are reliable. Not infrequently, some of the hypotheses tested are rejected; cf. Hall (1978, 1987) and Hansen and Singleton (1982, 1983). If they are, little can be said about the precise causes of rejection until a satisfactory empirical model <u>is</u> developed. Even if they are not rejected, the theoretical model commonly entails many "auxiliary" hypotheses for the empirical model which are testable (but often untested) and whose validity is <u>necessary</u> for the tests of the economic hypotheses to be reliable. Such auxiliary hypotheses include innovation disturbances, valid exogeneity assumptions, constant parameters, and the

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[1]The views expressed in this paper are solely the responsibility of the authors and should not be interpreted as reflecting those of the Board of Governors of the Federal Reserve System, the Banco Central de Venezuela, or other members of their staffs. The authors are indebted to Angus Deaton, David Hendry, Jeroen Kremers, Jaime Marquez, and Adrian Pagan for helpful discussions, and to Heidi Lyss for valuable research assistance. All numerical results were obtained using PC-GIVE Version 5.0; see Hendry (1987b).

ability to encompass (or account for) the results of other models. The last is necessary for any adequate empirical model: an inability to explain the properties of other models indicates the value of information contained in the other models, over and above that in the model being tested. Further, encompassing establishes an ordering across models such that an encompassing model serves as a sufficient statistic for all existing models; cf. Mizon and Richard (1986). Encompassing is particularly important when the alternative models have different economic and policy implications. In spite of that, encompassing is one of the least frequently used model evaluation criteria. However, encompassing proves crucial in this paper for excluding certain models from being empirically viable.

Many variables might be important in determining aggregate expenditure, including disposable income, liquidity, and inflation. The degree to which each plays a role and whether or not the relationship is stable over time have substantive implications for economic policy. In developed countries, liquidity and inflation appear to play significant roles, but ones very much secondary to that of income; cf. Hall (1987), and Hendry (1983) and Bean (1986) on US and UK evidence, respectively. Liquidity and inflation may be far more important in developing countries, given their typically more skewed income and wealth distributions, volatile rates of inflation, and incomplete capital markets. Venezuela is particularly well-suited for study: it has a consistently measured and reasonably defined set of data for the last two decades; and its economic activity has varied dramatically inter alia because of the debt crisis, a fundamental change in exchange rate policy, and its role as a major participant in OPEC. The importance of these factors is a theme recurring throughout the paper.

The paper is structured as follows. Section 2 presents a simple optimizing model in which agents aim to maintain constant expenditure-income and liquidity-income ratios but face adjustment costs. Section 3 sketches the econometric methodology, in which model design and evaluation play central roles. Section 4 describes the data. Based on the theoretical model of Section 2, Section 5 develops a parsimonious, statistically acceptable empirical model of annual consumers' expenditure on non-durables and services in Venezuela for 1970-85, and, in so doing, addresses issues of dynamic specification, parameter constancy, unit roots, cointegration, error-correction, and exogeneity. Disposable income, liquidity, and inflation determine expenditure in an economically sensible fashion, with each variable doing so through several channels. In spite of the turbulence of the Venezuelan economy over the sample, the empirical model is robust and has constant, well-determined parameter estimates. In fact, the volatility of the data is crucial for giving precise estimates with limited data. Section 6 considers other theoretical and empirical models via encompassing and as alternative interpretations of the empirical model. Following the encompassing approach, a large class of expectations-based and VAR models is found to be incompatible with the data. In particular, Hall's (1978) hypothesis (derived from the stochastic implications of the life cycle-permanent income hypothesis) that expenditure is a random walk and only predictable from its own past is rejected. The empirical model provides a sensible interpretation for why that is so. Further, the model's empirical properties lend some credibility to the economic and policy implications it entails. Section 7 discusses policy implications; Section 8 concludes. The paper may appear to devote much time to econometrics relative to that spent on economics, but it is necessary, both because the central econometric issues have economic counterparts and because, without proper econometric methodology, economic inference would be far more speculative in nature and lacking in empirical support.

2. The Economic Theory-Model

This section develops an economic agent's optimal rule for determining consumers' expenditure, given long-run targets for expenditure and assets and with short-run costs of adjustment. We assume ab initio that agents wish to have both expenditure and assets proportional to income in the long run. Those targets are embedded in a dynamic optimization problem with costs to being out of long-run equilibrium and to moving the control variable (expenditure) too rapidly. The agent's plan is the solution to that optimization, conditional on the agent's information set. Several common models are nested within that plan. Our framework generalizes upon that in Hendry and von Ungern-Sternberg (1981), and, in so doing, resolves a difficulty with their model. For the most part, we follow their notation. Starting from a multi-period utility-maximizing model with some agents facing liquidity constraints, Muellbauer and Bover (1986) derive both the proportionality targets and a similar agent plan. We prefer cost-minimization, noting the greater simplicity of derivation (c.f. Blundell (1988, pp. 24-26)) and the empirical difficulties with the expectations terms in Muellbauer and Bover's formulae (see Section 6 below).

Our theory-model is a partial one in that we assume at the outset that some prior steady-state maximization exercise has generated the equilibrium conditions:

$$W^e = K^* X \quad \text{and} \tag{2.1a}$$

$$A^e = B^* X \quad , \tag{2.1b}$$

where K^* and B^* are "constants", W_t is the agent's control variable, X_t is a forcing variable, A_t is an end-of-period stock defined as

$$A_t \equiv A_{t-1} + X_t - W_t \quad , \tag{2.2}$$

and a superscript ^{e} denotes steady-state equilibrium. For the empirical application below, W_t, X_t, and A_t may be interpreted as consumers' expenditure on non-durables and services, disposable income, and liquid assets.[2] However, they could represent the permanent or anticipated components, generating certain testable hypotheses if the process generating X_t changes within sample. Proportionality between (some form of) expenditure and income can be motivated by (e.g.) the permanent income or life-cycle hypotheses. Proportionality between assets and income may arise directly as an integral control mechanism (assets being the integral over past discrepancies between income and expenditure) or indirectly as a measure of liquidity constraints.[3] Although the two rationales have different policy implications, the empirical model will provide little evidence about which is a more appropriate interpretation for Venezuela.

In light of previous studies and the analytical properties of the corresponding models, we prefer to work with the logarithms of the data rather than their levels, necessitating the following steady-state approximation to (2.2):

$$\Delta a_t^e \approx H^*(x_t - w_t^e) \tag{2.3}$$

where $H^*=(1+g)/B^*$ and g is the steady state growth rate of X (and so of A and W). Throughout the paper, unless otherwise noted, capital letters denote both the generic name and the level; logs of scalars are in lower case. Defining the lag operator L as $Lz_t=z_{t-1}$ given an arbitrary series $\{z_t\}$, the difference operator Δ is (1-L); hence $\Delta z_t=(z_t-z_{t-1})$. More generally, $\Delta_r^q z_t=(1-L^r)^q z_t$. If q (or r) is undefined, it is taken to be unity. For convenience, equations (2.1a-b) can be re-written as:

$$w_t^e = k^* + x_t \quad \text{and} \tag{2.4a}$$

$$a_t^e = b^* + x_t \quad . \tag{2.4b}$$

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[2] Deaton (1972) derives an explicit role for wealth in the consumption function. See Salmon (1982), Nickell (1985), and Wilcox (1987) for more unified treatments of the optimization problem _per se_.

Our analysis is a partial one also because it omits the determination of expenditure on durables. Expenditure and income could incorporate services from and the depreciation of durables; see Patterson (1985).

[3] On the former, see _inter alia_ Phillips (1954, 1957), Stone (1966, 1973), and Hendry and von Ungern-Sternberg (1981); on the latter, see Pissarides (1978) and Muellbauer and Bover (1986).

Deviations from long-run targets may occur if outcomes are stochastic, so we posit a single-period loss function q_t for an agent's plans (a_t^p, w_t^p).

$$q_t = \lambda_1(a_t^p - a_t^e)^2 + \lambda_2(w_t^p - w_t^e)^2 + \lambda_3(w_t^p - w_{t-1})^2 - 2\lambda_4(w_t^p - w_{t-1})(w_t^e - w_{t-1}^e) - 2\lambda_5(w_t^p - w_{t-1})(w_t^e - w_{t-1}) \quad (2.5)$$

where $\lambda_i \geq 0$ $(i=1,\ldots,5)$. The first two terms on the RHS measure the costs associated with plans for assets and expenditure being out of line with their steady-state values. The third term assigns costs to changing expenditure and thereby will smooth the path of adjustment. Hendry and von Ungern-Sternberg (1981) use a loss function with the first four terms, giving this justification for the fourth.

> ..., when the primary objectives are to attain [(2.4a) and (2.4b)], it does not seem sensible to <u>quadratically</u> penalize changes in w_t^p when it is <u>known</u> that w_t^e has changed. Thus there is an offset term [the fourth] to allow <u>more adjustment at a given cost when w_t^e has changed than when it is constant</u>. (p. 240, italics in the original)

Noting (2.4a), Hendry and von Ungern-Sternberg use $-2\lambda_4(w_t^p - w_{t-1})\Delta x_t$ as the offset term, and it reduces the cost of changing w_t^p relative to w_{t-1} when the direction of change is the same as for the forcing variable x_t (or, equivalently, for w_t^e). Nickell (1985, p. 120) objects to such a term by the following illustration for which $\lambda_1=\lambda_5=0$. If an agent were out of long-run equilibrium last period but <u>now</u> $w_t^e=w_{t-1}\neq w_{t-1}^e$, then the agent (temporarily) moves <u>away</u> from w_t^e even though immobility implies being at the long-run equilibrium value. One solution is the following: the agent is concerned about moving w_t in the same direction that w_t^e has changed relative to the <u>realized</u> value w_{t-1} rather than the lagged equilibrium value w_{t-1}^e. One possible offset term is $-2\lambda_5(w_t^p - w_{t-1})(w_t^e - w_{t-1})$: it gives no incentive to change w_t if $w_t^e = w_{t-1}$ and so avoids Nickell's objection. For generality, we allow both λ_4 and λ_5 to be non-negative: below, we consider the effects on the agent's decision rule of restricting either (or both) to be zero.

Two equivalent decision rules can be obtained from minimizing the expected loss function with respect to planned expenditure, one in terms of equilibrium values and the other in terms of the forcing variable. Both offer insights, so we consider each in turn. The first is obtained by setting $[\partial E(q_t)/\partial w_t^p]$ equal to zero where the expectation $E(\cdot)$ is taken over some (for the time, arbitrary) information set T_t. Denoting such an expectation of a variable by a circumflex (^) over the variable, the decision rule is:

$$w_t^p - \hat{w}_{t-1} = (\lambda_4/\psi)(\hat{x}_t - \hat{x}_{t-1}) + \theta_2(\hat{w}_t^e - \hat{w}_{t-1}) + \theta_3(\hat{a}_{t-1} - \hat{a}_{t-1}^e) \quad (2.6)$$

where ψ and the θ_i's are functions of the λ_i's:

$$\psi = (H^{*2}\lambda_1 + \lambda_2 + \lambda_3) \quad (2.7a)$$

$$\theta_0 = [(\lambda_2+\lambda_5)k^* - \lambda_1 H^* b^*]/\psi \quad (2.7b)$$

$$\theta_1 = [H^*\lambda_1(H^*-1) + \lambda_2 + \lambda_4 + \lambda_5]/\psi \quad (2.7c)$$

$$\theta_2 = (H^{*2}\lambda_1 + \lambda_2 + \lambda_5)/\psi \quad (2.7d)$$

$$\theta_3 = (H^*\lambda_1)/\psi \quad . \quad (2.7e)$$

(θ_0 and θ_1 are used below.) All variables appear as expectations because the equilibrium values w_t^e and a_t^e are functions of x_t, and x_t may or may not be in T_t; and, depending upon the interpretation of x and w, their lagged values likewise may or may not be in T_t. The first term on the RHS of (2.6) results from Hendry and von Ungern-Sternberg's offset term. The expression $(\hat{w}_t^e - \hat{w}_{t-1})$ in the second term is the discrepancy between the expected equilibrium value of w_t and (the expectation of) w_{t-1}, the most recent realized value of w_t. The expression $(\hat{a}_{t-1} - \hat{a}_{t-1}^e)$ in the final term is the extent to which the (expected) realized value of assets at the <u>beginning</u> of the current period is out of line with their expected equilibrium value. Thus, the latter two terms in (2.6) measure the optimal adjustments to w_t because of long-run disequilibria in expenditure and asset holdings. The first term captures adjustment due to changes in the equilibrium value of w_t.[4] Substituting (2.4a) and (2.4b) into (2.6) and re-arranging gives

$$w_t^p - \hat{w}_{t-1} = \theta_0 + \theta_1(\hat{x}_t - \hat{x}_{t-1}) + \theta_2(\hat{x}_{t-1} - \hat{w}_{t-1}) + \theta_3(\hat{a}_{t-1} - \hat{x}_{t-1}) \ , \qquad (2.8)$$

which is the decision rule expressed in terms of the forcing variable x.

The realized value w_t may be stochastic and not necessarily equal to w_t^p. We allow for that discrepancy by positing

$$w_t = w_t^p + u_t \qquad\qquad u_t \sim NI(0,\sigma_u^2) \qquad (2.9)$$

where u_t is independent of T_t and so of w_t^p.[5] Agents make plans which are optimal conditional on the information available to them, and discrepancies between those plans and actual outcomes arise solely because of shocks outside of (and orthogonal to) that information. This efficient use of information corresponds to the statistical concept of innovation errors entailed by conditioning; that ties economic theory to statistical properties of a corresponding empirical model. Realized expenditures are:

$$w_t - \hat{w}_{t-1} =$$
$$\theta_0 + \theta_1(\hat{x}_t - \hat{x}_{t-1}) + \theta_2(\hat{x}_{t-1} - \hat{w}_{t-1}) + \theta_3(\hat{a}_{t-1} - \hat{x}_{t-1}) + u_t \ . \qquad (2.10)$$

In a non-stochastic steady-state with $\Delta x = g$, (2.3) implies a solution of

$$(W/X) = D \cdot (A/X)^{\phi} \qquad (2.11)$$

where $\phi = \theta_3/\theta_2 > 0$ and $D = \exp\{[\theta_0 - (1-\theta_1)g]/\theta_2\}$, linking the steady-state conditions (2.1a) and (2.1b).

Although the loss function q_t appears simple, several well-known models (including partial adjustment and error-correction) exist as special cases to the corresponding optimal rule for w_t^p in (2.6)-(2.8). When only quadratic terms appear in (2.6) (i.e., $\lambda_4 = \lambda_5 = 0$), then $\theta_1 + \theta_3 = \theta_2$ and (2.10)

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[4] If $\lambda_4 = 0$, that term vanishes and (2.6) is a partial-adjustment equation involving long-run targets. However, when equilibrium values in (2.6) are expressed in terms of the forcing variable x, neither that new equation (nor its estimable companion) need be partial adjustment, even when $\lambda_4 = 0$.

[5] The disturbance u_t also embodies the approximation error in (2.3) and mis-specification error. The bearing of such errors on empirical model design is addressed in Section 3.

simplifies to eliminate $(\hat{x}_t-\hat{x}_{t-1})$, giving a natural generalization of partial adjustment (but here, adjustment to disequilibria in both x and a).

$$w_t - \hat{w}_{t-1} = \theta_0 + \theta_2(\hat{x}_t-\hat{w}_{t-1}) + \theta_3(\hat{a}_{t-1}-\hat{x}_t) + u_t \quad . \qquad (2.12)$$

If the asset-to-income target is also unimportant (i.e., $\lambda_1=0$), the "pure" partial adjustment model is obtained. Alternatively, if only λ_1 is zero, then $\theta_3=0$ in (2.10) to give an error-correction mechanism with no integral control:

$$w_t - \hat{w}_{t-1} = \theta_0 + \theta_1(\hat{x}_t-\hat{x}_{t-1}) + \theta_2(\hat{x}_{t-1}-\hat{w}_{t-1}) + u_t \quad ; \qquad (2.13)$$

cf. Davidson, Hendry, Srba, and Yeo (1978). Finally, if the agent's information set T_t is (X_t^o, W_{t-1}^o) where $Z_j^o=(z_0,\ldots,z_{j-1},z_j)$ for arbitrary z, then all the expectations in (2.10), (2.12), and (2.13) are actual values.

The choice of the information set, the measures of expenditure, income, and assets, and aggregation over agents and time are issues important in linking this theoretical model to an empirical model of expenditure, but a thorough treatment of them is outside this paper's scope. Even so, econometric methodology can help understanding these and other facets of modeling, so it is discussed in the following section.

3. Econometric Methodology

This section discusses the interplay between economic theory and the data, and so the role of testing and design in empirical modeling. Modeling itself is viewed as an attempt to characterize data properties in simple parametric relationships which remain reasonably constant over time, account for the findings of pre-existing models, and are interpretable in the light of the subject matter (e.g., theory, institutions, historical background). Hence, econometric models are viewed as simplifications of the underlying real-world process generating the data; test statistics serve as evaluation and design criteria for these models.[6]

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[6] For general expositions and bibliographies, see Hendry and Richard (1982, 1983), Hendry (1983, 1987a), Hendry and Wallis (1984), Ericsson and Hendry (1985), and Spanos (1986).

The economic theory upon which an empirical model is based often contains many ceteris paribus conditions which may not hold in fact. Thus, one of the tasks of the econometrician is to choose an empirical model which both embodies the economic theory (and so permits interpretation of the model itself) and allows for the presence of any significant factors not fully specified by the economic theory. If an empirical model fails to allow for the presence of such factors, then inference, forecasting, and policy analysis based upon it are likely to be unreliable. That argues for initially specifying a rather general empirical model.

Even so, any empirical model (including such a "general" specification) is inherently a reduced re-parameterization of the data generation process and is obtained by implicitly performing two types of operations on that process: integrating out (i.e., excluding) variables and treating one set of variables as being determined by contemporaneous values of another. Those two operations correspond to the statistical concepts of marginalizing and conditioning respectively; the former includes aggregation over agents and time, and the latter inter alia reflects the notion of economic agents forming plans contingent on (weakly exogenous) contemporaneous variables. Such operations necessarily entail that the "error" is a derived rather than an autonomous process, containing everything not modeled explicitly. Likewise, the parameters of the empirical model are derived, and are defined by the model's specification and the orthogonality conditions of the estimation technique. Hence, it may be possible to design an empirical model to satisfy both data-based and theory criteria.

To illustrate, consider the standard linear regression model:

$$y_t = \beta' x_t + u_t \qquad t=1,\ldots,T \qquad (3.1)$$

where x_t are the k relevant regressors and β is the (presumably constant) parameter of economic interest. For (3.1) as an empirical model, both β and $\{u_t\}$ are derived, not autonomous. For instance, for OLS, β is defined by the orthogonality condition $\{E[x_t(y_t-\beta' x_t)]=0;\ t=1,\ldots,T\}$ and u_t by $y_t-E(y_t|x_t)$, so the properties of both may vary according to the choice and functional form of x_t.[7] For reliable inference, it is important to evaluate whether the empirical model (3.1) satisfies the properties which its proponents claim (e.g., white noise, innovation disturbances, constant parameters). Further, it may be possible to design empirical models which have those properties and so may permit valid inference.

How well or poorly designed an empirical economic model is depends upon its ability to capture salient features of the data and to deliver reliable inference on economic issues (e.g., coefficient estimates, predictions, policy). Hence, we place considerable emphasis on a battery of test statistics which serve as criteria both for evaluating existing specifications and for designing new ones. The criteria are standard and relate to goodness-of-fit, absence of residual autocorrelation and heteroscedasticity, valid exogeneity, predictive ability, parameter constancy, the statistical and economic interpretation of estimated coefficients, the validity of *a priori* restrictions, and the ability of a model to account for properties of alternative models. Table 1 summarizes the corresponding statistics, which are arranged by the types of information generating testable null hypotheses: the data of one's own model, the measurement system of the data, economic theory, and the data of alternative models. Most of the model's empirical properties bear on its economic as well as statistical interpretation: four deserve particular note.

Parameter constancy is at the heart of model design, both from a statistical and from an economic perspective. Most estimation techniques require parameter constancy for valid inference, and those that seem not to do so, still posit "meta-parameters" assumed constant over time. Since economic systems are far from being constant, and the coefficients of

[7] Similar formulae exist for more general models and estimation techniques.

Table 1. Evaluation/Design Criteria

Information Set	Null Hypothesis	Alternative Hypothesis	Sources
(A) own model's data			
(A1) relative past	innovation errors	first-order residual autocorrelation	Durbin and Watson (1950, 1951)
"	"	2: q^{th}-order residual autocorrelation	Box and Pierce (1970); Godfrey (1978), Harvey (1981, p. 173)
"	"	3: invalid parameter restrictions	Johnston (1963, p. 126)
"	"	4: q^{th}-order ARCH	Engle (1982)
"	"	5: heteroscedasticity quadratic in regressors	White (1980, p. 825), Nicholls and Pagan (1983)
"	"	6: q^{th}-order RESET	Ramsey (1969)
"	normality of the errors	7: skewness (SK) and excess kurtosis (EK)	Jarque and Bera (1980)
(A2) relative present	weakly exogenous regressors	8: invalid conditioning	Sargan (1958, 1980), Engle, Hendry, and Richard (1983)
(A3) relative future	constant parameters, adequate forecasts	1: parameter non-constancy, predictive failure	Fisher (1922), Chow (1960), Brown, Durbin, and Evans (1975), Hendry (1979)
(B) measurement system	data admissibility	"impossible" predictions of observables	-
(C) economic theory	theory consistency	"implausible" coefficients, predictions	-
(D) alternative models' data			
(D1) relative past	variance dominance	relative poor fit	Hendry and Richard (1982)
"	variance encompassing	inexplicable observed error variance	Cox (1961, 1962), Pesaran (1974)
"	parameter encompassing	significant additional variables	Johnston (1963, p. 126), Mizon and Richard (1986)
(D2) relative present	exogeneity encompassing	inexplicable valid conditioning	Hendry (1988)
(D3) relative future	forecast encompassing	informative forecasts from rival model	Chong and Hendry (1986)

Note: The numbers preceding the alternative hypotheses label the statistics in Sections 5 and 6.

derived ("non-structural" or "reduced form") equations may alter when any of the underlying parameters or data correlations change, it is important to identify empirical models which have reasonably constant parameters and which remain interpretable when some change occurs. That puts a premium on good theory. Conversely, empirical models with constant parameterizations in spite of "structural change" elsewhere in the economy may provide the seeds of fruitful research in economic theory.

Recursive estimation of an equation provides an incisive tool for investigating parameter constancy, both through the sequence of estimated coefficient values and via the associated Chow statistics for constancy.[8] Again, using the standard linear regression model to illustrate, the OLS estimator of β in (3.1) is $\hat{\beta}_T$ where:

$$\hat{\beta}_t = (X_t'X_t)^{-1}X_t'Y_t \tag{3.2}$$

with $X_t = (x_1 \ldots x_t)'$ and $Y_t = (y_1 \ldots y_t)'$. It is relatively simple and computationally inexpensive to obtain the <u>entire sequence</u> of OLS estimates $\{\hat{\beta}_t: t=h,\ldots,T\}$ starting with some initial number of observations h ($h \geq k$). Recursive estimation can generate voluminous output: T-h+1 estimates of each coefficient, a standard error for each estimate, the sequence of estimated innovations (recursive residuals), and the many possible sequences of Chow statistics. Graphs can portray this information in an effective, concise manner, and we will use them to do so below. The Chow statistics also play crucial roles (a) for testing weak exogeneity indirectly through testing the conjunction of hypotheses embodied in super exogeneity and (relatedly) (b) for testing feedback versus feedforward empirical models; cf. Engle, Hendry, and Richard (1983) and Hendry (1988).

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[8] These tests of constancy are intimately related to tests of forecast accuracy; cf. Brown, Durbin, and Evans (1975), Hendry (1979), and Kiviet (1986). Dufour (1982) elegantly summarizes recursive techniques and their implications.

Cointegration links the economic notion of a long-run relationship between a set of variables with statistical models of those variables.[9] For example, in the theory-model of Section 2, the realized values of W and A need not be proportional to X, but economic forces (agent optimization) bring them back into line if they are not so. With shocks in every period, proportionality never need hold, yet agents work to absorb those shocks and W and A never drift "too far" from their equilibrium values. If x_t is non-stationary (e.g., integrated of order one, I(1)), w_t and a_t also will be so, but some linear combination of w_t, x_t, and a_t may be stationary because of agents' behavior, i.e., making that set of variables co-integrated. The choice of normalization of the cointegrating vector is an unresolved issue, but both economics and the data can help. Theory may suggest which variables agents aim to control and on which ones they condition their plans, and parameter constancy in an empirical model is not invariant to normalization when the economy exhibits structural change. Cointegration implies the existence of an error-correction representation of the relevant variables, leading to the third issue, dynamic specification. Normalization and conditioning lead to the fourth issue, exogeneity.

Dynamic specification influences model design because of the statistical and economic importance of white-noise, innovation disturbances. Dynamics may appear in empirical models because the agent's optimization explicitly dictates its presence (as in Section 2 and Nickell (1985)), because agent behavior implies cointegrated variables and hence dynamics, because *ceteris paribus* conditions of the theory-model may not hold in fact, or any combination thereof. As a rule, dynamic mis-specification invalidates inference, so dynamics cannot be safely ignored, and a general

[9] See Engle and Granger (1987) and Hendry (1986). See Engle (1987) and Dolado and Jenkinson (1987) for recent surveys on cointegration.

specification (at the outset, at least) often is advisable.[10]

Weak exogeneity can occur when agents condition on information. If they use it efficiently, innovation errors are implied, raising the issue of dynamic specification. Weak exogeneity is testable, often as an implication of super exogeneity (and so of models having constant parameters). Section 5 further discusses exogeneity in the context of the empirical model.[11]

The econometric approach above contrasts with those in which the data serve only for estimating parameters of the theory-model. Estimation is an important issue, but, in itself, provides little guidance on the value or lack thereof of the empirical model obtained. An empirical model is unlikely to allow reliable statistical or economic inference, forecasting, or policy analysis unless it is "well-designed" in the sense that it does not violate either the assumptions made at the outset or the numerous testable implications of those assumptions.

4. The Data Series and Data Transformations

In light of the modeling approach outlined in Section 3, a satisfactory empirical model must account for basic features of the data, and a model may be designed to do so. Conversely, an empirical model failing to account for them may be re-designed to do so better. Instead of being "data mining", looking at the data can be invaluable in model construction. Further, entire classes of models may be incompatible with observed data properties; nb. Deaton (1987). The remainder of this section describes the data.

The data are annual values for Venezuela for 1968-1985 of consumers' expenditure on non-durables and services (C), national disposable income

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[10] See Hendry, Pagan, and Sargan (1984) and Trivedi (1986) for extensive discussion on theoretical aspects of dynamic specification.

[11] The four distinct concepts of exogeneity, namely weak, strong, super, and strict, correspond to different notions of being "determined outside the model under consideration" according to the purposes of the inferences being conducted, i.e., conditional inference, prediction, policy analysis, and forecasting, respectively. For detailed discussion of these concepts, see Engle, Hendry, and Richard (1983) who build upon those in Koopmans (1950).

(Y), and end-of-period liquidity (L) as measured by the monetary aggregate M_2.[12] The unit of currency is the Bolívar (abbreviated Bs), and its value was pegged at 4-4.5 Bs/US$ until February 1983, after which time substantial devaluation occurred under a multi-tier exchange rate system. These series are real (1968 Bolívares) per capita; the price deflator (P) is for all consumers' expenditure. Appendix A gives details of the data and sources.

Before examining the data in detail, we note several caveats. We chose to model the determinants of consumers' expenditure on only non-durables and services so as to measure relatively homogeneous expenditure and to avoid difficulties associated with calculating the flow of services from durables. The expenditure data allow finer disaggregation, presenting an avenue for further research; cf. Palma (1976). Alternative measures of liquidity are available or feasibly constructed: an adjustment for Venezuelan holdings of liquid assets in foreign countries may improve the measure substantially although the mechanism by which such assets influence <u>domestic</u> expenditure is unclear. Price deflators are available for aggregate consumers' expenditure and for its major components. Most of those series are similar, and empirically it mattered little which one was used. That is in part a consequence of the analytical properties of the theory-models and the models estimated: in the long-run, they are invariant as to whether expenditure, income, and liquidity are nominal or real, total or per capita. Because population growth is quite constant throughout the sample, the empirical results are virtually the same whether the data are aggregate or per capita. However, the population size increased by 70% from 1968 to 1985, so per capita data are more amenable to <u>economic</u> interpretation.

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[12] We use the notation (C, Y, L) to distinguish from the (potentially latent) economic-theoretic variables (W, X, A) in Section 2.

Earlier data are available, but involve numerous and substantial re-definitions of the series, rendering them less informative. More frequent data are available, but only for some series.

Measurement of income raises numerous issues. Personal disposable income is clearly preferable to national disposable income, but data for the former are not available and our attempts to construct such a series have been unfruitful. Neither measure accounts for the sometimes substantial "inflation tax" from holding liquid assets. In an attempt to improve that situation, we constructed an adjusted series Y_t^* defined as $Y_t-(\Delta p_t)L_{t-1}$, which nets current income of the financial loss from holding last period's end-of-period assets in the presence of current inflation and so aims to better capture "that accrual which would leave real wealth intact"; cf. Hendry and von Ungern-Sternberg (1981).[13] To provide some sense of the magnitudes involved, for 1973 (L_{t-1}/Y_t) and Δp_t are .27 and 5.5% respectively, so $Y_t^*=[1-\Delta p_t(L_{t-1}/Y_t)]\cdot Y_t=.985Y_t$; but by 1984 (L_{t-1}/Y_t) and Δp_t are .52 and 16.8%, with Y_t^* only $.91Y_t$. Adjusted and measured income differ substantially in the latter instance.

Consider some basic properties of the data themselves. From Figure 1, three distinct episodes are evident in the behavior of expenditure and income. Before 1974, both grew at relatively moderate rates. Because of dramatically increased petroleum revenues, income increased by over 30% in 1974 and, through 1981, remained on a plateau at 155-170% of the level of 1968 income. From 1974, expenditure grew rapidly (but less so than income), leveling off in the late 1970's, with the expenditure to income ratio in 1981 being virtually the same as in 1968. From 1981 to 1985, (real per capita) income plummeted at 7% per annum, but expenditure remained relatively constant. Figure 2 graphs the growth rates of expenditure and income: the levels of the series are highly correlated but not their growth rates, and the volatility of income is notably greater than that of

[13] Also, Hendry and von Ungern-Sternberg (1981) include a (possibly non-unit) coefficient on $-(\Delta p_t)L_{t-1}$ to allow for scale effects in the mis-measurement of Δp and L. We would like to do so as well, but given the very limited number of observations and the value and extensive use of recursive estimation, we set aside that generalization for further research.

Figure 1. The logs of real per capita expenditure (c) and of real per capita income (y^*) (with means adjusted).

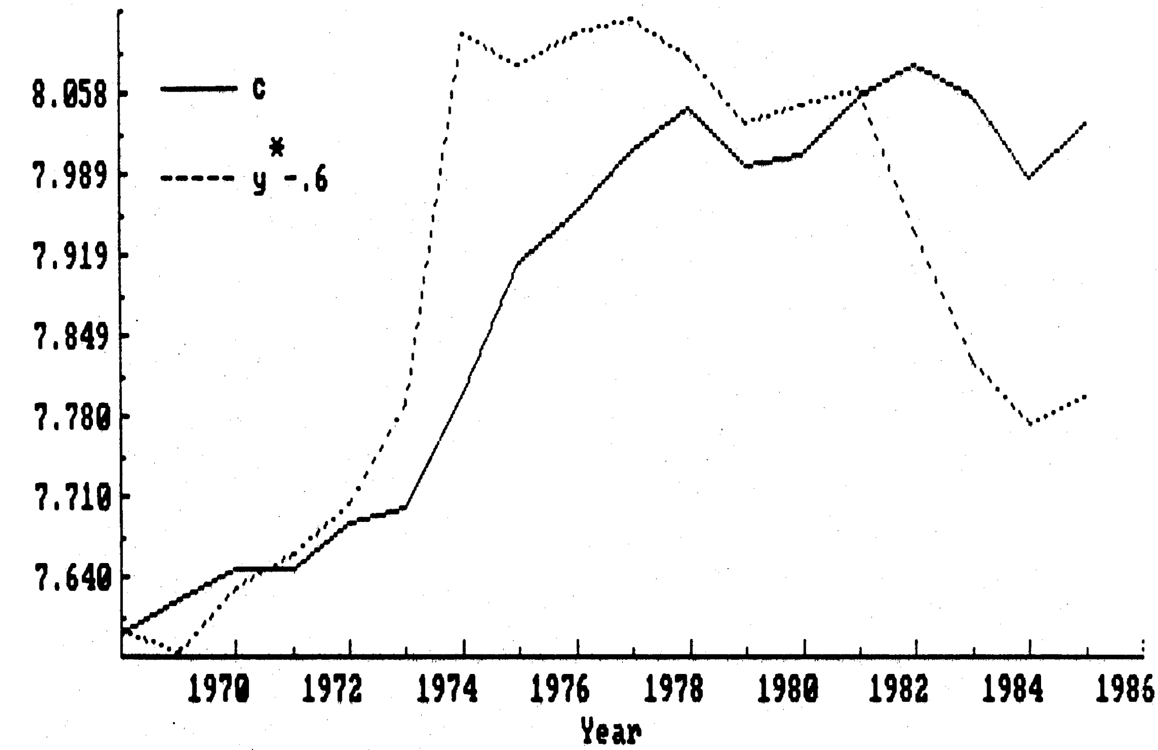
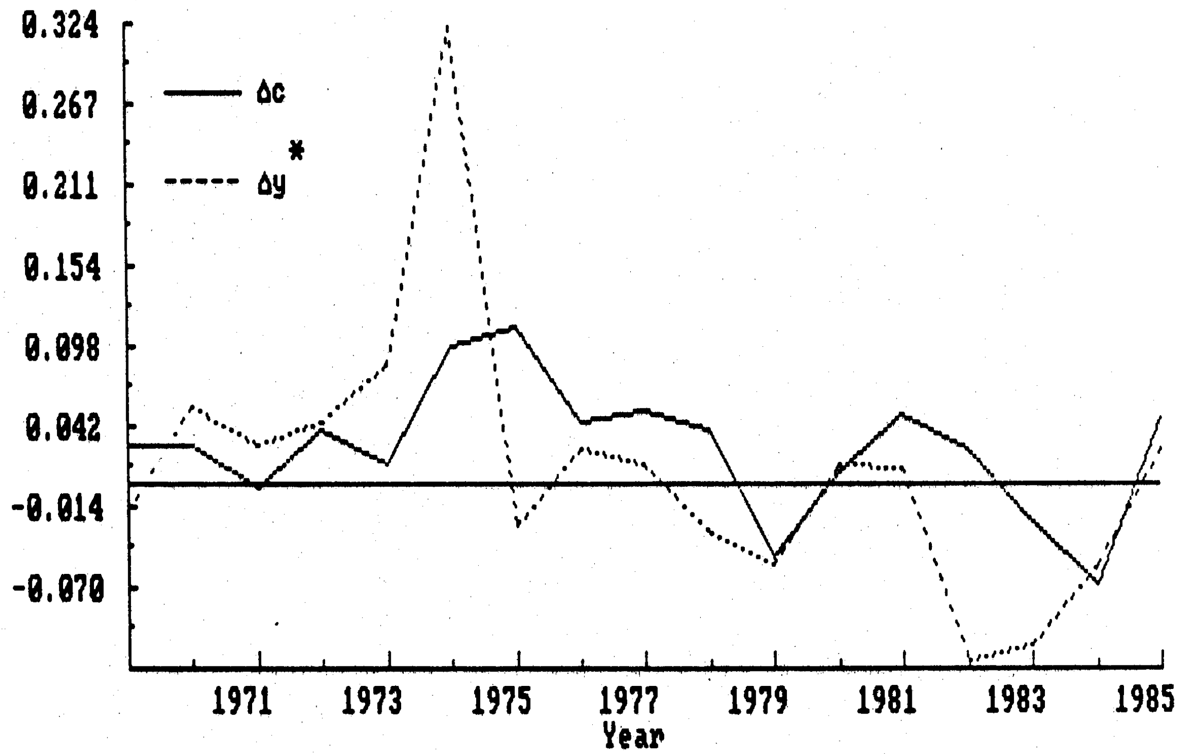


Figure 2. The growth rates of real per capita expenditure (Δc) and of real per capita income (Δy^*).



expenditure. A static relationship between c and y is unlikely to generate such properties, and even a dynamic bivariate one may not.

One possible explanation for the differing responses of expenditure to income is the dramatic change in liquidity: Figure 3 graphs the (logs of) the expenditure-income and liquidity-income ratios. Those ratios rise by 27% and 135% over the period. Consistent with the theory in Section 2, the substantial increase of the latter ratio could account for the stability of expenditure in the 1980's in spite of income falling. Both ratios markedly fall in 1974 and rise in 1982-83: large changes in income are primarily responsible rather than changes in expenditure or liquidity; cf. Figure 1.

Digressing briefly, petroleum revenues are central to understanding the income process in Venezuela. Venezuela is a member of OPEC, with petroleum exports equaling approximately a quarter of GNP and accounting for roughly 90% of the value of all its exports, much of which is sold to the United States.[14] As Figure 4 shows, the nominal growth rates of national disposable income and petroleum exports (denoted Δyn and Δxp) are highly correlated and, in fact, are virtually indistinguishable after accounting for their different scales. This, and the relationship between US and Venezuelan inflation noted below, play a crucial part in Section 6 when we attempt to distinguish between feedback (conditional) models as developed in Section 2 and feedforward (expectations-based) models.

Finally, consider prices. Figure 5 compares US inflation (denoted Δp^+) and Venezuelan inflation (Δp). The two were similar through 1978, but, while the United States experienced moderately higher inflation after the second oil-price shock in 1979, Venezuelan inflation jumped to nearly 20%, falling during the world-wide recession in the early 1980's and sharply increasing again in 1984. The higher inflation rates during the second half

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[14] Conversely, a large proportion of imports by Venezuela is from the United States.

Figure 3. The logs of the consumption-income ratio ($c-y^*$) and of the liquidity-income ratio ($\ell-y^*$).

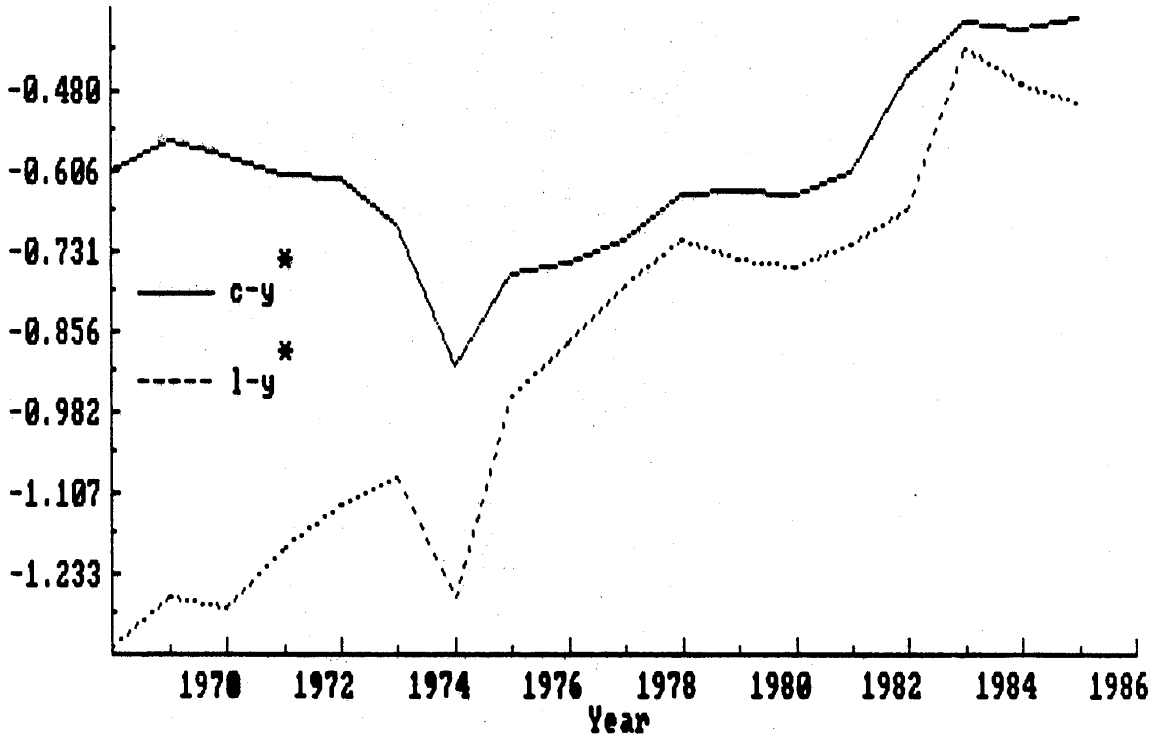


Figure 4. The growth rates of nominal income (Δyn) and of nominal petroleum exports (Δxp) (with ranges adjusted).

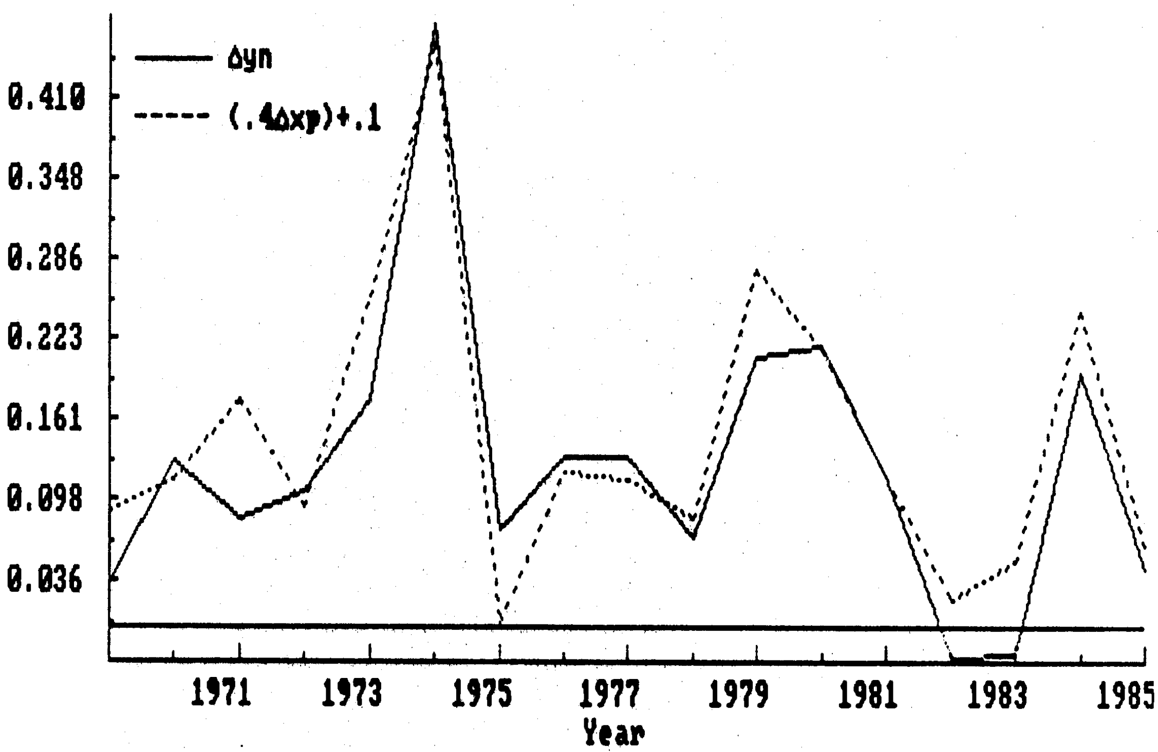


Figure 5. Venezuelan inflation (Δp) and US inflation (Δp^{+}).

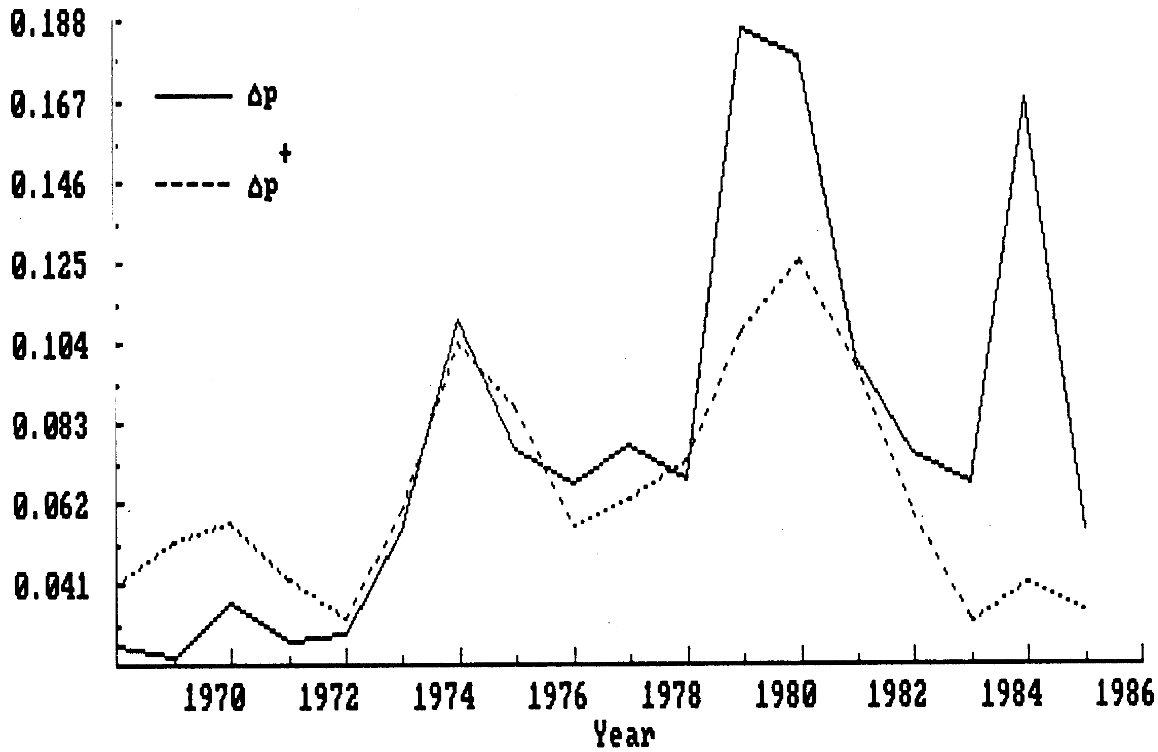
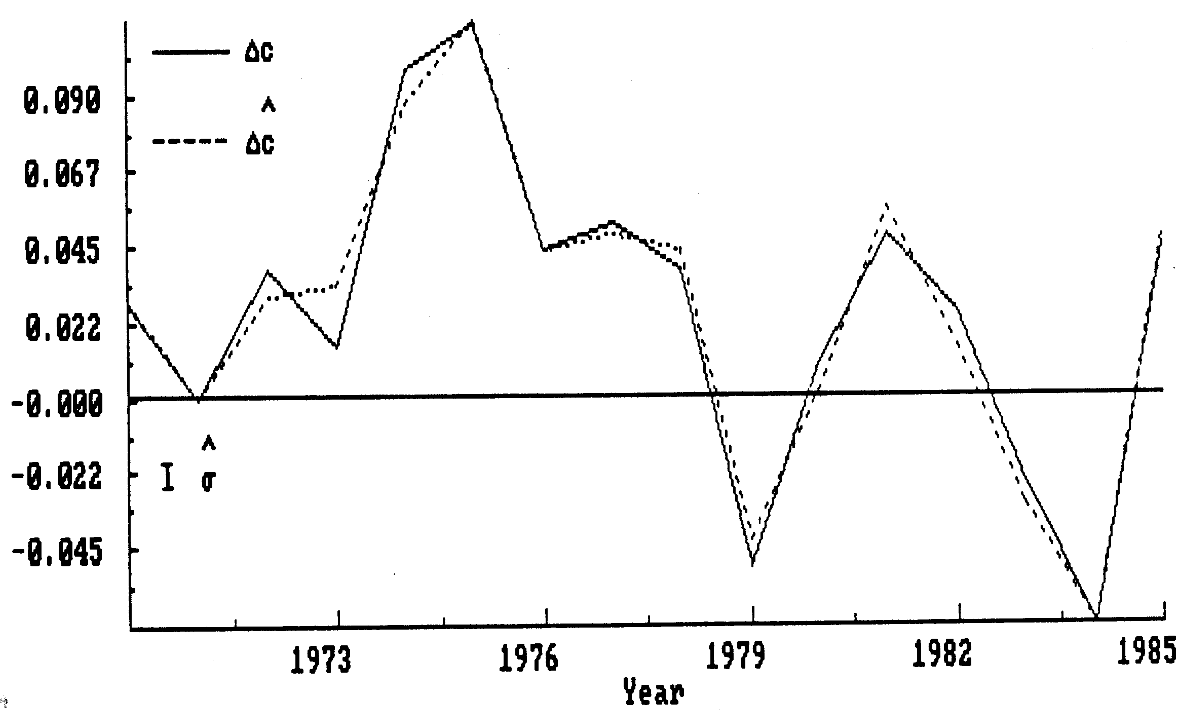


Figure 6. Equation (5.4): actual and fitted values for Δc.



of the sample, combined with greater liquidity, led to the more substantial discrepancies between Y and Y^* observed above. Inflation may also affect expenditure directly, e.g., if agents have difficulty discerning relative from absolute price changes; cf. Deaton (1977). Further, the theory-model in Section 2 takes inflation as given, yet the constants of proportionality B^* and K^* might vary as inflation does. And, without specifying the precise timing of information availability and agents' actions, the existing theory can not exclude <u>short</u>-run inhomogeneity of consumers' expenditure with respect to prices. All three reasons imply inflation and its lags entering (2.10) additively.

Before moving to empirical model specification and testing, one more issue must be addressed: the number of observations. There are eighteen in total, with only sixteen used in estimation. In many contexts, that number would be small and, given the relatively large number of coefficients to be estimated, impracticably so. The absolute number of observations does restrict the complexity of the empirical model, placing a premium on both good economic theory and effective econometric methodology. However, the variability of the data as well as the number of observations determines the precision of the resulting estimates and the power of tests. For instance, the variance of the OLS estimator for (3.1) is approximately $\sigma_u^2 \cdot [T \cdot E(x_t x_t')]^{-1}$ for stationary data. Although T is small, movements in data such as $c-y^*$ and $\ell-y^*$ are large: fluctuations in comparable post-war US data are an <u>order</u> of magnitude smaller. Such properties help account for the small estimated standard errors obtained below.[15]

5. <u>An Empirical Model of Consumers' Expenditure</u>

This section develops a conditional econometric model of Venezuelan consumers' expenditure on non-durables and services. It is interpretable in

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[15] Large fluctuations in the data also can reduce problems of measurement error.

the framework of the agent optimization problem of Section 2, with income, liquidity, and inflation all helping determine expenditure. However, since significant, economically plausible estimates are not sufficient for reliable inferences, the remainder of the section evaluates the empirical model by the criteria discussed in Section 3, focusing on dynamic specification, parameter constancy, cointegration, and exogeneity. Encompassing is addressed in Section 6.

In order to establish a baseline innovation variance, consider the following general autoregressive distributed lag relationship for consumers' expenditure conditional upon income, prices, and liquidity.

$$c_t = \sum_{i=1}^{2} (\alpha_i^* c_{t-i} + \delta_i^* \ell_{t-i}) + \sum_{i=0}^{2} (\beta_i^* y^*_{t-i} + \gamma_i^* p_{t-i}) + \mu_0 + \mu_1 D_{01} + u_t \quad (5.1)$$

D_{01} is a +1/-1 dummy for 1970-71 to account for apparent measurement errors in consumers' expenditure for those years; see Appendix B. With loss of generality, we assume that $(c-y^*)$ and $(\ell-y^*)$ are cointegrated, implying that $1-\Sigma(\alpha_i^*+\delta_i^*)=\Sigma\beta_i^*$ and that prices enter only as inflation (i.e., $\Sigma\gamma_i^*=0$).[16] With cointegration, (5.1) can be re-written as in (5.2) below with all coefficients in (5.2) unrestricted. The assumptions imbedded in (5.1) are testable and will be tested (in part) below, but the tests available lack power because of the small sample size: even with the cointegration restrictions imposed, (5.1) implies ten unrestricted coefficients estimated with sixteen observations.

$$\begin{aligned} \Delta c_t = {} & \alpha_1 \Delta c_{t-1} + \sum_{i=0}^{1} (\beta_i \Delta y^*_{t-i} + \gamma_i \Delta p_{t-i}) + \delta_1 \Delta \ell_{t-1} + \alpha_0 (c-y^*)_{t-1} \\ & + \delta_0 (\ell - y^*)_{t-1} + \mu_0 + \mu_1 D_{01} + u_t \end{aligned} \quad (5.2)$$

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[16] The economic theory of Section 2 implies cointegration of $(c-y^*)$ and $(\ell-y^*)$, provided agents can maintain a finite variance on their realized errors.

This equation has the general form of the theoretical model (2.10), but contains additional lags because the theory is vague about the precise dynamic structure. The lagged rate of change of liquidity is included because the theory-model does not exclude short-run inhomogeneity of consumers' expenditure with respect to liquidity (as with prices).

Data coherence and dynamic specification. Table 2 lists the coefficient estimates for (5.2), the associated conventional and heteroscedasticity-consistent standard errors (denoted $(\cdot)$ and $[\cdot]$ respectively; see White (1980) and Nicholls and Pagan (1983)), and diagnostic statistics, to the extent that the latter are available for such an over-parameterized model. $\xi_i(q)$ and $\eta_i(q,r)$ denote statistics which have central $\chi^2(q)$ and $F(q,r)$ distributions respectively under a common null and against the i^{th} alternative, where i is as listed on Table 1.

The model in Table 2 has two notable features. First, the residual standard error is slightly above 1%: any new model will require a similar or smaller equation standard error as a necessary condition for encompassing the model in Table 2. Second, three economically sensible and statistically acceptable classes of parametric restrictions are apparent. The lagged rates of change of expenditure and of liquidity are insignificant: a single lag on each level is sufficient to capture those parts of the dynamics. The coefficients on the current and lagged growth rates of income are equal: that can be interpreted as a statistical smoothing of income in order to extract changes which are more permanent. Changes in the growth rate of income also immediately affect the budget constraint and liquidity, so giving the coefficient on current income (or its growth rate) several interpretations. The coefficient on current inflation is approximately twice the magnitude of that on lagged inflation, and opposite in sign. Imposing that restriction implies the term $\Delta p_t + \Delta^2 p_t$ (with negative coefficient), which is a predictor of next period's inflation, optimal if

Table 2. A general autoregressive-distributed lag representation for real per capita consumers' expenditure conditional on incomes, prices, and liquidity.

Variable	lag i: 0		lag i: 1	
Δc_{t-i}	-1.0	(-) [-]	-.081	(.163) [.126]
Δy^*_{t-i}	.224	(.045) [.054]	.211	(.056) [.030]
Δp_{t-i}	-.553	(.086) [.062]	.260	(.118) [.092]
$\Delta \ell_{t-i}$	-		.004	(.057) [.042]
$(c-y^*)_{t-i}$	-		-.169	(.062) [.035]
$(\ell-y^*)_{t-i}$	-		.070	(.027) [.017]
D_{01}	.026	(.009) [.001]	-	
Constant	.006	(.033) [.027]	-	

T = 1970-1985 $\quad R^2 = .9743 \quad \hat{\sigma} = 1.199\% \quad dw = 2.59$

$\xi_2(2) = 3.3 \quad \eta_6(1,5) = .5 \quad \xi_7(2) = .2$

Notes:

1. The dependent variable is Δc_t. Even so, the equation is in levels, not in differences, noting the transformation of variables from the re-parameterization of (5.1) to obtain (5.2).

2. Because of the small number of degrees of freedom remaining, there are few diagnostic statistics. More statistics appear with the restricted equations (5.3) and (5.4), in which cases the validity of those statistics is conditional upon the simplifying restrictions.

prices vary quadratically. Economically, that might capture a smaller desired D in (2.11) in the face of higher inflation, or the aim of consumers (e.g.) to save <u>more</u> now in anticipation of higher inflation so as to be able to consume more closely the same amount in real terms when that higher inflation arrives.[17] Re-estimating (5.2) with those restrictions imposed gives (5.3).

$$\begin{aligned} \widehat{\Delta c}_t = {} & \underset{\substack{(.047)\\ [.051]}}{.460}(\Delta_2 y_t^*/2) - \underset{\substack{(.028)\\ [.024]}}{.270}(\Delta p+\Delta^2 p)_t - \underset{\substack{(.023)\\ [.016]}}{.141}(c-y^*)_{t-1} \\ & + \underset{\substack{(.013)\\ [.010]}}{.072}(\ell-y^*)_{t-1} + \underset{\substack{(.016)\\ [.011]}}{.021} + \underset{\substack{(.007)\\ [.001]}}{.026}D_{01} \end{aligned} \qquad (5.3)$$

$T = 1970\text{-}1985 \quad R^2 = .9725 \quad \hat{\sigma} = .9603\% \quad dw = 3.03 \quad \xi_2(3) = 6.5$

$\eta_3(4,6) = .1 \quad \eta_6(2,8) = .4 \quad \xi_7(2) = .5$

The F-statistic η_3 tests the restrictions which obtain (5.3) from Table 2, and it is insignificant.

Surprisingly, the estimates in (5.3) are remarkably close to those obtained by Hendry and von Ungern-Sternberg (1981) with quarterly data for the United Kingdom. Comparable estimates to those for the growth rate of

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[17] The presence of the term $\Delta p+\Delta^2 p$ (and the interpretation of it in the text) may be viewed in several complementary ways. First, current purchases of consumer <u>durables</u> are one way of "saving" now for consumption later. When a higher inflation is anticipated, the budget constraint implies lower current expenditure on non-durables and services to do so. Second, because the income distribution is highly skewed in Venezuela, increasing liquidity by saving may be one of the few means available to a substantial fraction of the population for avoiding substantial negative shocks to real consumption in the face of anticipated higher inflation. That applies particularly to periods like 1978-79 and 1982-84 during which real per capita income fell: see Figure 2. Third, consumers may have difficulty distinguishing between relative and aggregate changes in prices, so inflation may induce saving; see Deaton (1977). Finally, $\Delta p+\Delta^2 p$ might capture the degree to which the coefficient on $-(\Delta p_t)L_{t-1}$ in Y^* is different from unity, cf. Hendry and von Ungern-Sternberg (1981, p. 247).

The term $\Delta p+\Delta^2 p$ does <u>not</u> appear to be a proxy for the interest rate, nor does the interest rate appear important in determining expenditure. The rate offered at commercial banks is statistically insignificant if added to (5.4) below, and that result is insensitive as to whether the rate is contemporaneous and/or lagged. Further, $\Delta p+\Delta^2 p$ remains highly significant with the interest rate included in (5.4).

income and the lags of $(c-y^*)$ and $(\ell-y^*)$ are .50, -.16, and .072.[18] The implied long-run elasticity ϕ of the expenditure-income ratio with respect to the liquidity-income ratio is .51, remarkably close to one-half and to Hendry and von Ungern-Sternberg's estimate of .44. Because of that, and in an effort to design as parsimonious a model as possible, we re-estimated (5.3) with ϕ=.5.

$$\hat{\Delta c}_t = \underset{\substack{(.030)\\ [.033]}}{.457(\Delta_2 y^*_t/2)} - \underset{\substack{(.026)\\ [.023]}}{.270(\Delta p+\Delta^2 p)_t} - \underset{\substack{(.019)\\ [.014]}}{.142[(c-y^*)-.5(\ell-y^*)]_{t-1}}$$

$$+ \underset{\substack{(.004)\\ [.004]}}{.019} + \underset{\substack{(.007)\\ [.001]}}{.026D_{01}} \tag{5.4}$$

$T = 1970\text{-}1985 \quad R^2 = .9725 \quad \hat{\sigma} = .9160\% \quad dw = 3.05 \quad \xi_2(3) = 6.7$

$\eta_{3_a}(5,6) = .1 \quad \eta_{3_b}(1,10) = .01 \quad \eta_4(1,9) = .5 \quad \eta_5(8,2) = .3$

$\eta_6(2,9) = .2 \quad \xi_7(2) = .6$

The two F-statistics η_{3_a} and η_{3_b} test the restrictions which obtain (5.4) from Table 2 and (5.3) respectively, and, in conjunction with the other statistics, they suggest the statistical acceptability of those restrictions on the information available.[19]

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[18] Even so, the time series and data moments of the two countries differ markedly.

[19] The power of these tests is quite possibly very low, although the few degrees of freedom involved are balanced to some extent by the considerable information content of each observation; cf. Section 4. Even so, the Durbin-Watson and Box-Pierce statistics provide some weak evidence of <u>negative</u> autocorrelation in the residuals. (Other tests for autocorrelation are not calculated by PC-GIVE for such a small sample.) Additional measurement errors of the sort described for 1970-71 in Appendix B or dynamic mis-specification could be responsible. In support of the latter, the Durbin-Watson statistic in (5.4) falls to 2.27 when the one- <u>and</u> two-period lags on Δy^*_t and Δp_t are added; but those lags are insignificant, whether tested jointly or individually. Additional data may resolve this issue. We also note that, in their empirical model of consumers' expenditure, Hendry and von Ungern-Sternberg (1981) detect negative autocorrelation of approximately the same magnitude as that implied by the Durbin-Watson statistic in (5.4). This <u>class</u> of model may imply negatively autocorrelated residuals, either due to the finite sample properties of the estimation method used or from a type of mis-specification characteristic to error-correction models with integral control; but that remains conjecture.

The coefficients in (5.4) have sensible values and highlight the roles of contemporaneous and lagged information. In spite of income, liquidity, and prices being highly inter-correlated, the RHS variables in (5.4) are virtually orthogonal to each other, giving an economically plausible partitioning of the information set. All the economic variables may be interpreted as affecting consumers' liquidity, and so consumers' inclination (or ability) to spend. Income growth and inflation affect liquidity both immediately (by increasing current liquidity directly and through the inflation tax) and in the future (because of their long-run effects and because of the interpretation of $\Delta p+\Delta^2 p$ as a predictor of inflation). The median lag for expenditure responding to changes in income is short, under a year; but adjustment is gradual thereafter. Figure 6 shows the actual and fitted values for the rate of change in real per capita expenditure (Δc) over the sample period: $\hat{\sigma}$ is less than 1%, but c̲ changes by over 4% for half the observations and by over 10% for two observations. That implies a remarkable fit relative to the observed data fluctuations.

The steady-state solution implied by (5.4) is:

$$(C/Y^{*}) = (L/Y^{*})^{.5} \cdot \exp(.135 - 3.82g - 1.90\Delta p) \qquad (5.5)$$

where g and Δp are the steady-state growth rates of income and prices. Equations (5.4) and (5.5) correspond to (2.10) and (2.11) of the theory-model. The exponential in (5.5) is the coefficient of proportionality D in (2.11), and is a function not only of g (as postulated) but of Δp (which belonged to the host of *ceteris paribus* assumptions in the theory-model). The coefficients in the exponential imply that agents spend more relative to income for a given liquidity ratio when income is growing slower and/or inflation is lower. The latter implies improved *future* liquidity. Thus, although (5.4) is a conditional model, its particular specification gives it and its long-run solution (5.5) a forward-looking characteristic.

Equation (5.4) has a plausible economic interpretation in light of (2.10) and (2.11), is a statistically acceptable simplification of Table 2 (given our information set), and appears adequately dynamically specified, but it must satisfy many additional conditions to be a satisfactory empirical model. These include: the estimated parameters are constant; expenditure, income, and liquidity are cointegrated; income and prices are weakly exogenous; and (5.4) encompasses all existing models. We now turn to the first three.

Parameter constancy. To investigate the constancy of the model (5.4), we adopt the recursive estimator described in Section 3. Figure 7 records the sequences of one-step residuals (i.e., $y_t - x_t'\hat{\beta}_t$) and corresponding calculated equation standard errors ($\pm 2\hat{\sigma}_t$ is shown centered on zero). The residual standard error has varied little, in spite of the dramatic fluctuations in the series themselves. The estimated coefficients also are constant: Figures 8, 9, and 10 show the sequences of estimates for the coefficients on $(\Delta_2 y_t^*/2)$, $(\Delta p + \Delta^2 p)_t$, and $[(c-y^*)-.5(\ell-y^*)]_{t-1}$, together with plus-or-minus twice their sequentially estimated standard errors which provide an approximate 95% confidence interval. The accrual of information over time is particularly apparent for inflation, with its coefficient estimated standard error falling by two-thirds in 1979. That is not so surprising, given the marked change in the properties of inflation in 1979; cf. Figure 5. Finally, none of the Chow statistics for the sequences (a) {1979, 1980, ..., 1985}, (b) {1979, 1979-80, ..., 1979-85}, or (c) {1979-85, 1980-85, ..., 1985} is significant at the 5% level.[20]

[20] In fact, no individual Chow statistic of any sequence exceeded even a third of its corresponding 5% critical value. These Chow statistics (and the other diagnostic statistics) served as design criteria, so their insignificance in part relfects our ability to design an adequate empirical model. However, (5.4) is virtually identical to Campos's (1984) final specification using data for 1968-82 and upon which this study is based. Thus, forecasts for 1983-85 present a veritable test of the empirical model (5.4) because the data for those years were not available to Campos (1984).

Figure 7. Equation (5.4): one-step residuals and the corresponding calculated standard errors.

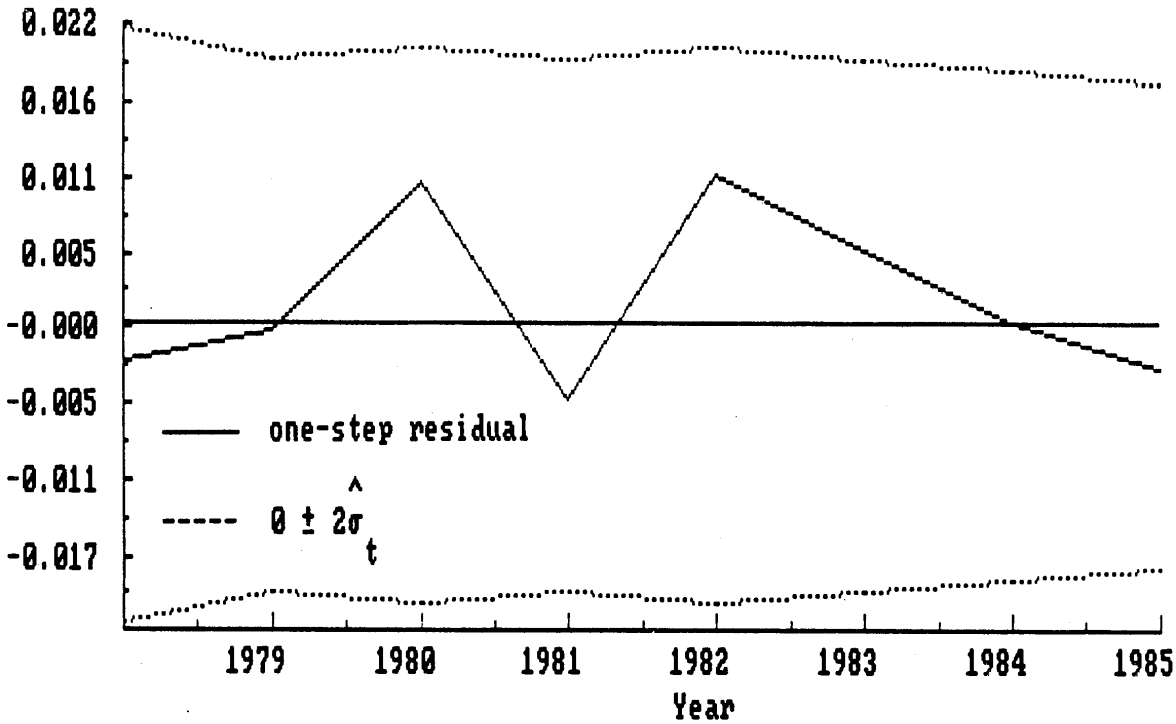


Figure 8. Equation (5.4): recursive estimation of the coefficient on the two-year average growth rate of income ($\Delta_2 y_t^*/2$).

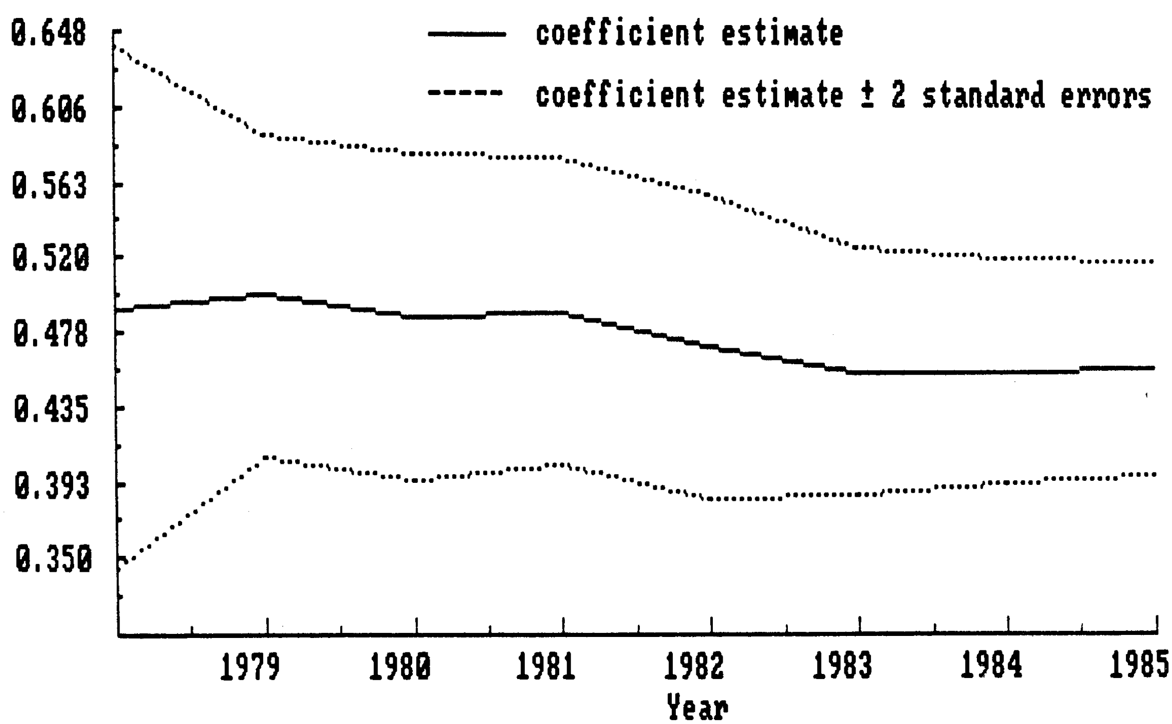


Figure 9. Equation (5.4): recursive estimation of the coefficient on the inflation rate term $(\Delta p+\Delta\Delta p)_t$.

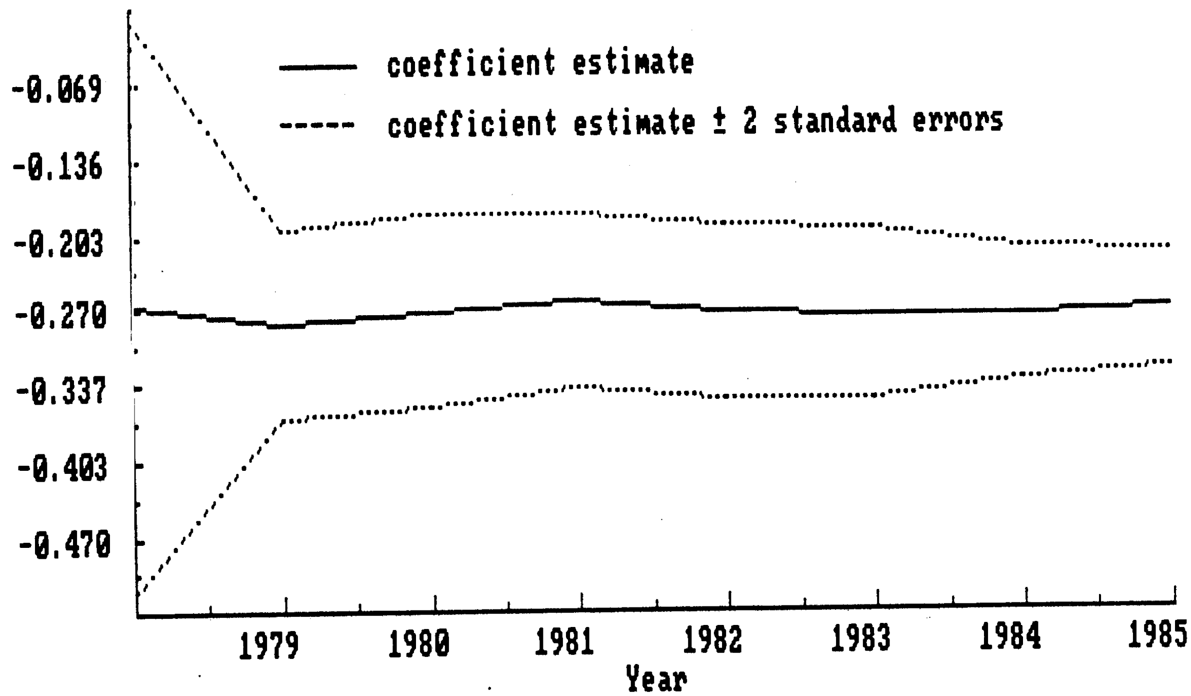
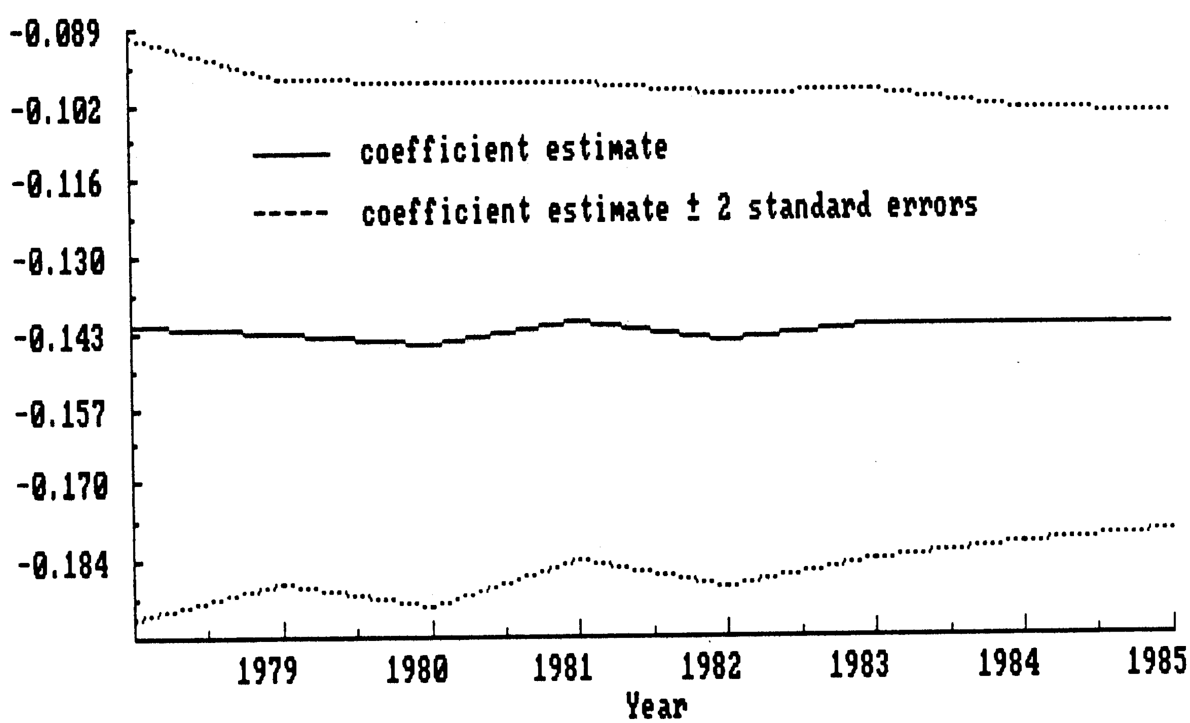


Figure 10. Equation (5.4): recursive estimation of the coefficient on the error-correction term $[(c-y^{*})-.5(\ell-y^{*})_{t-1}]$.



To evaluate further the numerical and statistical accuracy of (5.4), we estimated it using the data for 1970-80 only and calculated forecasts for the remaining five years using the sub-sample coefficient estimates but actual values of the RHS variables. As Figures 8-10 show, the sub-sample estimates are virtually those in (5.4) but have somewhat larger estimated standard errors. Figure 11 plots the actual and forecast values of Δc for 1981-85 with 95% confidence intervals for each forecast. Figure 12 plots the actual and fitted values of Δc through 1980, with actual and forecast values thereafter. Actual values lie well within the confidence intervals, and the fit of the model (as measured by $\hat{\sigma}$) improves slightly with the additional data. In brief, the conditional formulation (5.4) both fits well and is remarkably constant over a turbulent decade and a half of the Venezuelan economy.

Cointegration. In testing for the existence of cointegration, it is necessary to establish first that the individual series are I(1) and then that there exists some non-trivial function of them which is I(0). The most common approach is to test for a unit root in the univariate representations of the individual series and then in the least-squares linear combination. Table 3 lists the values of the Sargan-Bhargava (Durbin-Watson) statistic and the augmented Dickey-Fuller statistic (ADF) for the logs of expenditure, income, liquidity, prices, and the expenditure-income and liquidity-income ratios. No tests reject the hypothesis of a unit root at the 5% level although the power of these tests is low for a sample size of sixteen and a root close to unity (e.g., .8).[21] The Engle-Granger static cointegration regression for expenditure, income, and liquidity is:

[21] Cf. Evans and Savin (1981), Dickey and Fuller (1979, 1981), and Sargan and Bhargava (1983).

Note also that I(0) series with structural breaks can look like I(1) series: for Venezuela, the former may more appropriately describe the data.

Figure 11. Equation (5.4) estimated over 1970-80: actual and forecast values of Δc for 1981-85, and 95% confidence bands for the forecasts.

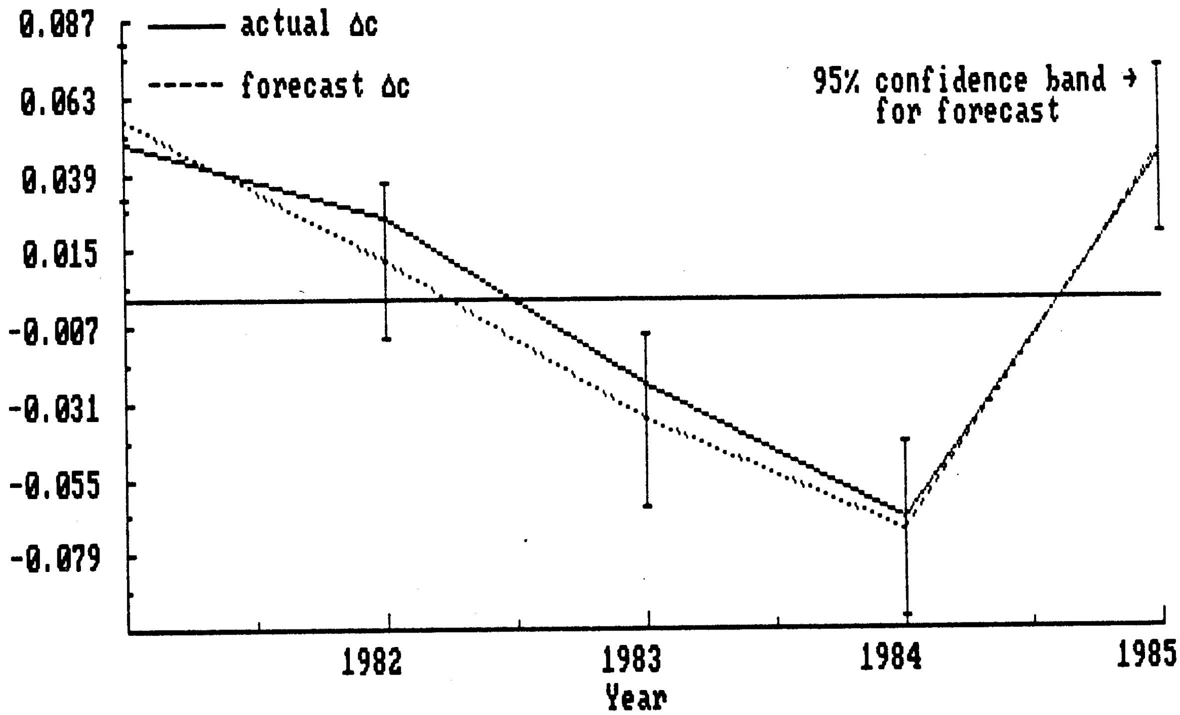


Figure 12. Equation (5.4) estimated over 1970-80: actual, fitted, and forecast values of Δc.

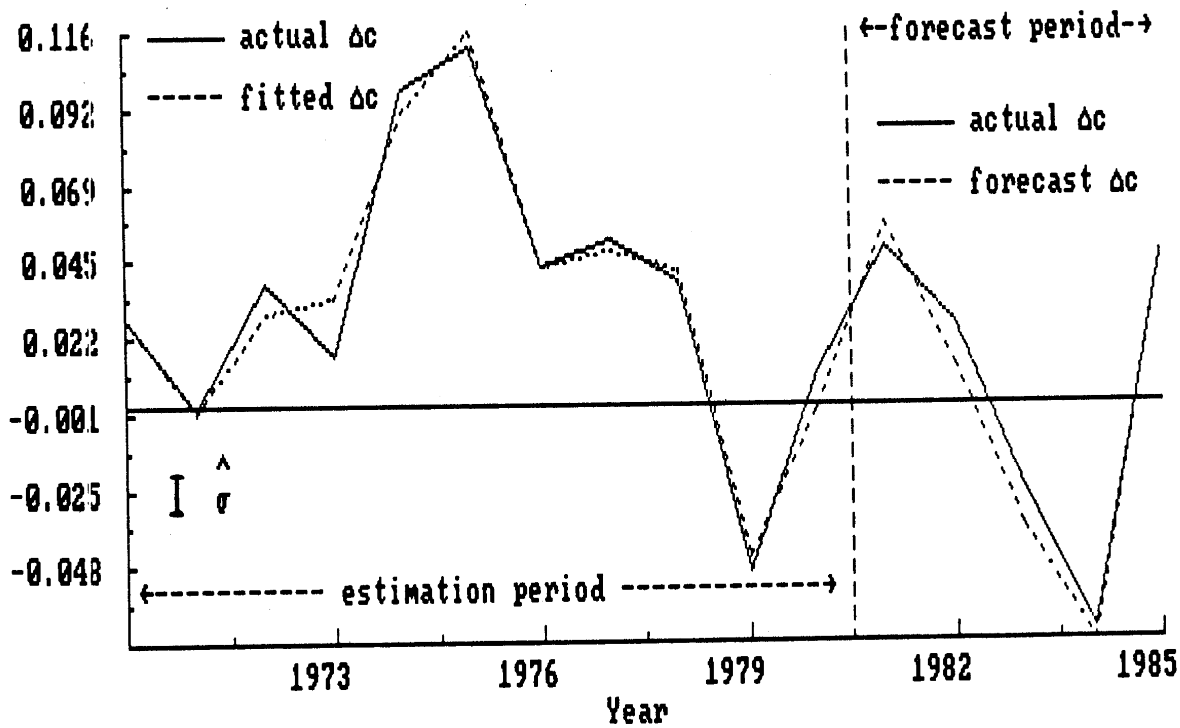


Table 3. Univariate statistics for testing unit roots.

Variable	dw	ADF(0)	ADF(1)
c	.08	-1.52	-1.42
y^*	.23	-1.52	-1.97
ℓ	.07	-2.05	-2.01
p	.04	1.93	.70
$c-y^*$	.33	-.55	-.68
ℓ_1-y^*	.13	-.62	-.48
$\hat{u}_{(5.6)}$	2.01	-3.89	-3.26
$\hat{u}_{(5.7)}$	.37	-1.28	-1.43
ecm^*	.23	-1.72	-2.00

Notes:

1. The variables $\hat{u}_{(5.6)}$ and $\hat{u}_{(5.7)}$ are the residuals from equations (5.6) and (5.7).

2. The variable ecm^* is $(c-y^*)_t-.5(\ell_{t-1}-y_t^*)$. Lagged ℓ is used here and in (5.6) and (5.7) because it measures end-of-period liquidity.

3. The Durbin-Watson (dw), Dickey-Fuller (ADF(0)), and augmented Dickey-Fuller (ADF(1)) statistics are calculated over the maximum number of observations available: 1968-85, 1969-85, and 1970-85, respectively.

$$\hat{c}_t = .082y^*_t + .36\ell_{t-1} + 4.45 \qquad (5.6)$$

T = 1968-1985 $R^2 = .9658$ $\hat{\sigma} = 3.549\%$ dw = 2.01 ADF(0) = -3.89

ADF(1) = -3.26 .

The Sargan-Bhargava statistic is at the upper end of the inconclusive region (albeit offering no evidence against white-noise residuals as well), and the Dickey-Fuller statistics are significant at about the 5% level. However, the theory-model in Section 2 is in terms of the expenditure-income and liquidity-income ratios, suggesting a second cointegration regression.

$$(\widehat{c-y^*})_t = .26(\ell_{t-1}-y^*_t) - .36 \qquad (5.7)$$

T = 1968-1985 $R^2 = .4077$ $\hat{\sigma} = 11.36\%$ dw = .37 ADF(0) = -1.28

ADF(1) = -1.43

None of the test statistics in (5.7) provides much evidence in favor of cointegration although their power to do so is minimal.

One alternative to these univariate methods is to estimate the conditional error-correction mechanism for c_t and test the significance of the error-correction term therein. Monte Carlo results in Banerjee, Dolado, Hendry, and Smith (1986) suggest that the size of the test based on the t-ratio of the error-correction term is approximately correct for the 5% level and that this test is much more powerful than the Sargan-Bhargava test. That follows because the Sargan-Bhargava and Dickey-Fuller statistics for (e.g.) (5.7) impose short-run as well as long-run homogeneity on $(c-y^*)$, and cointegrated series need not be short-run homogeneous. Those univariate tests remain consistent even with short-run inhomogeneity, but they can lose power. Conversely, the error-correction model allows for short-run inhomogeneity, so tests using the t-ratio of the error-correction term may be more powerful. To wit, the highly significant estimate on $(\Delta_2 y^*_t/2)$ in (5.4) implies substantial short-run inhomogeneity of expenditure with respect to income, and the t-ratio on the error-correction term is -7.6.

Further, Banerjee et al. (1986) show that, in finite samples, the static estimator of the cointegrating vector can be badly biased towards zero relative to the corresponding estimator in the error-correction model, particularly when R^2 is small in the static regression. In (5.7), R^2 is only .41, and the static and dynamic estimates of the cointegrating vector are .26 and .51, respectively, in line with those finite sample results. Thus, although these results on cointegration are qualified, the disparities between the static and dynamic regressions are in line with existing finite sample evidence for highly dynamic cointegrated variables and argues for full dynamic modeling, both to achieve greater power and to reduce bias.[22]

Exogeneity. One important, testable assumption is the weak exogeneity of income and prices, i.e., the validity of conditioning upon them for statistical inference (estimation and testing). We test for weak exogeneity both by estimating (5.4) by instrumental variables and comparing those results with (5.4), and by constructing marginal models for income and inflation which are non-constant over time, from which it follows that income and inflation are super exogenous (and hence weakly exogenous) for the parameters in (5.4).

Given the marked dependence of the Venezuelan economy on the United States for manufactured goods and on petroleum exports for the purchase of those goods, we selected Δp^{+}_{t-1}, Δp^{+}_{t-2}, Δxp_{t-1}, and Δc_{t-1} as instruments for $(\Delta_2 y^{*}_{t}/2)$ and $(\Delta p+\Delta^2 p)_t$. The results are in (5.8).

$$\begin{aligned}\Delta\tilde{c}_t = {} & \underset{(.053)}{.425}(\Delta_2 y^{*}_{t}/2) - \underset{(.080)}{.353}(\Delta p+\Delta^2 p)_t - \underset{(.029)}{.155}[(c-y^{*})-.5(\ell-y^{*})]_{t-1} \\ & + \underset{(.007)}{.025} + \underset{(.009)}{.027}D_{01}\end{aligned} \tag{5.8}$$

$T = 1970\text{-}1985 \quad \tilde{\sigma} = 1.287\% \quad \xi_2(3) = 5.1 \quad \eta_4(1,9) = .2$

$\eta_5(8,2) = .5 \quad \xi_7(2) = .2 \quad \xi_8(2) = .4$

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[22] If (5.4) is estimated by Engle and Granger's two-step procedure in which the residuals from (5.7) replace the error-correction term in (5.4), very similar parameter estimates result, but $\hat{\sigma}$ increases to 1.152%.

Sargan's (1958) statistic ξ_8 does not reject the validity of the instruments. Additionally, the estimated coefficients are negligibly different from those in (5.4) and the estimated standard errors are somewhat larger, as would be expected if (5.4) were well-specified with no simultaneity bias. Similar results obtain for other instrumental variables.

Super exogeneity requires both weak exogeneity and structural invariance, so finding super exogeneity implies weak exogeneity. Demonstrating super exogeneity relies on showing that the parameters of the conditional process remain constant even though the marginal process changes. The precise justification of this test for super exogeneity requires a subtle argument, cf. Richard (1980), Engle et al. (1983), Engle and Hendry (1985), and Hendry (1988). Even so, the nature of the argument is well-illustrated with the following simple expectations-based model in which x_t and y_t are simultaneously determined. This model also lays the foundation for testing feedback versus feedforward models in Section 6 and refines the links between parameter constancy and super exogeneity.

Suppose that the correct specification for y_t is in terms of the expectation of x_t given z_t, rather than x_t itself, e.g.,

$$y_t = \beta^{*\prime}E(x_t|z_t) + u_t^* \tag{5.9}$$

with

$$x_t = E(x_t|z_t) + v_t \tag{5.10}$$

where $E(z_t u_t^*)=0$ and $E(x_t|z_t)\neq 0$ by assumption, $E(z_t v_t)=0$ by construction, and β^* (not necessarily equal to β in (3.1)) is the parameter of interest.[23] By substitution,

$$y_t = \beta^{*\prime}x_t + (u_t^* - \beta^{*\prime}v_t) \quad . \tag{5.11}$$

[23]The variables x and u here are not necessarily equivalent to the variables in Section 2 designated by the same letters.

Instrumental variables estimation of (5.11) using the instrument z_t is consistent for β^* because z_t is uncorrelated with the composite error term. However, the OLS estimator for (5.11) generally is not consistent for β^*:

$$\bar{\beta} \quad \equiv \quad \text{plim } \hat{\beta}^* \quad = \quad \beta^* + [E(x_t x_t')]^{-1}[E(v_t u_t^*) - E(v_t v_t')\beta^*] \qquad (5.12)$$

from the standard formulae for simultaneity bias; see Bronfenbrenner (1953). For ease of illustration, suppose that x_t is generated as an AR(1):

$$x_t \quad = \quad \rho_0 x_{t-1} + v_t \qquad v_t \sim NI(0,\sigma_0^2) \qquad t=1,\ldots T_0 \qquad (5.13a)$$

$$x_t \quad = \quad \rho_1 x_{t-1} + v_t \qquad v_t \sim NI(0,\sigma_1^2) \qquad t=T_0+1,\ldots,T \qquad (5.13b)$$

where x_t is univariate and ρ_0 may not equal ρ_1. If $\rho_0=\rho_1\equiv\rho$ and $E(v_t u_t^*)=0$, (5.12) simplifies to

$$\bar{\beta} \quad = \quad \beta^* - (1-\rho^2)\beta^* \qquad (5.14)$$

for $|\rho|<1$ and $\rho\neq0$. Consistent estimation of β^* in (5.9) requires information additional to x_t itself (e.g., knowing that z_t enters (5.10)). Efficient estimation requires modeling the process for x_t in full. That is, x_t is not sufficient information for conducting inference about β^*: hence x_t cannot be weakly exogenous for β^*.[24] Conversely, $\hat{x}_t$ $(\equiv E(x_t|z_t))$ may be weakly exogenous for β^*; but $\hat{x}_t$ is unobservable, which necessitates <u>knowing</u> something about the process for x_t in order to conduct inference about β^*.

In terms of "textbook econometrics", OLS estimation of (5.11) is inconsistent for the underlying constant parameter of interest β^* and, instead, plim $\hat{\beta}^*$ is a function of both β^* and the parameters of the process for x_t, e.g., $\bar{\beta} \equiv \text{plim } \hat{\beta}^* = \bar{\beta}(\beta^*,\rho)$ when x_t is AR(1). Hence, the OLS estimate will vary as the parameters of the marginal process do: the Lucas critique applies. Conversely, establishing that $\rho_0\neq\rho_1$ (or more generally, that the parameters of the process for x_t change) and that β in (3.1) is nevertheless constant over [1,T] shows that x_t is super exogenous for β.

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[24] This is true even though the marginal process for x_t may involve lags of x_t and no other variables. Strict exogeneity of x_t is neither sufficient nor necessary for its weak exogeneity.

Technically speaking, in that case, the parameters of interest β are invariant to those changes in the marginal process for x_t. The constancy of the parameters in (5.4) has already been established; it remains to show that marginal processes for prices and income are non-constant.

The marginal processes for inflation and income are simple: Venezuelan prices depend upon US prices and income depends upon petroleum exports. As follows immediately from Section 6 below, it is <u>not</u> necessary to test the constancy of fully specified marginal processes: it suffices that <u>a</u> marginal process be non-constant. Estimating over 1968-78, we obtain:

$$\widehat{(p-p^+)}_t = \underset{\substack{(.093)\\ [.109]}}{.681}(p-p^+)_{t-1} - \underset{\substack{(.35)\\ [.41]}}{1.21} \qquad (5.15)$$

T = 1968-1978 $\quad R^2 = .8563 \quad \hat{\sigma} = .9533\% \quad dw = 1.66 \quad \eta_1(7,9) = 33.7$

$\eta_3(2,7) = 1.0 \quad \eta_5(2,6) = .4 \quad \eta_6(1,8) = 5.2 \quad \xi_7(2) = 1.0$.

The model appears reasonably well-specified within sample, with neither short-run nor long-run homogeneity between Venezuelan and US prices rejected.[25] However, the model exhibits substantial predictive failure over the remainder of the sample, as is evident from the one-step residuals and their standard errors in Figure 13 and from η_1 in (5.15). If Venezuelan inflation is modeled as an AR(1) process, similar results obtain.

For <u>nominal</u> income, the marginal process is the error-correction model:

$$\widehat{\Delta yn}_t = \underset{\substack{(.033)\\ [.029]}}{.134}\Delta xp_{t-1} + \underset{\substack{(.026)\\ [.017]}}{.407}\Delta(xp_t - yn_{t-1}) + \underset{\substack{(.067)\\ [.048]}}{.284}(xp-yn)_{t-1} + \underset{\substack{(.088)\\ [.065]}}{.476} \qquad (5.16)$$

T = 1970-1985 $\quad R^2 = .9581 \quad \hat{\sigma} = 2.618\% \quad dw = 1.14 \quad \eta_{1_a}(1,7) = 8.2$

$\eta_{1_b}(1,7) = 8.8 \quad \eta_3(2,10) = 1.6 \quad \eta_4(1,10) = .5 \quad \xi_7(2) = .4$.

In (5.16), income has a long-run elasticity of unity with respect to petroleum exports and a median lag of under a year. However, the parameters

[25] Note that (5.15) is equivalent to the model in which p_t equals p_t^+ plus a constant and error, and the error is AR(1).

Figure 13. Equation (5.15): one-step residuals and the corresponding calculated standard errors.

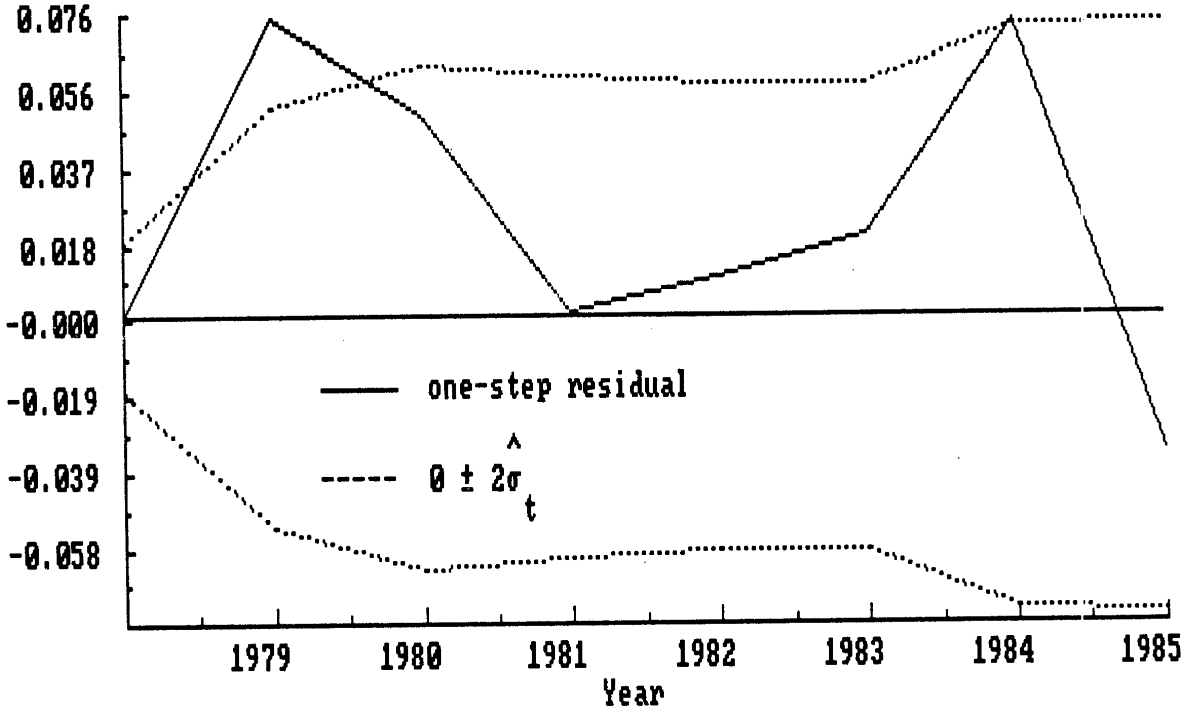
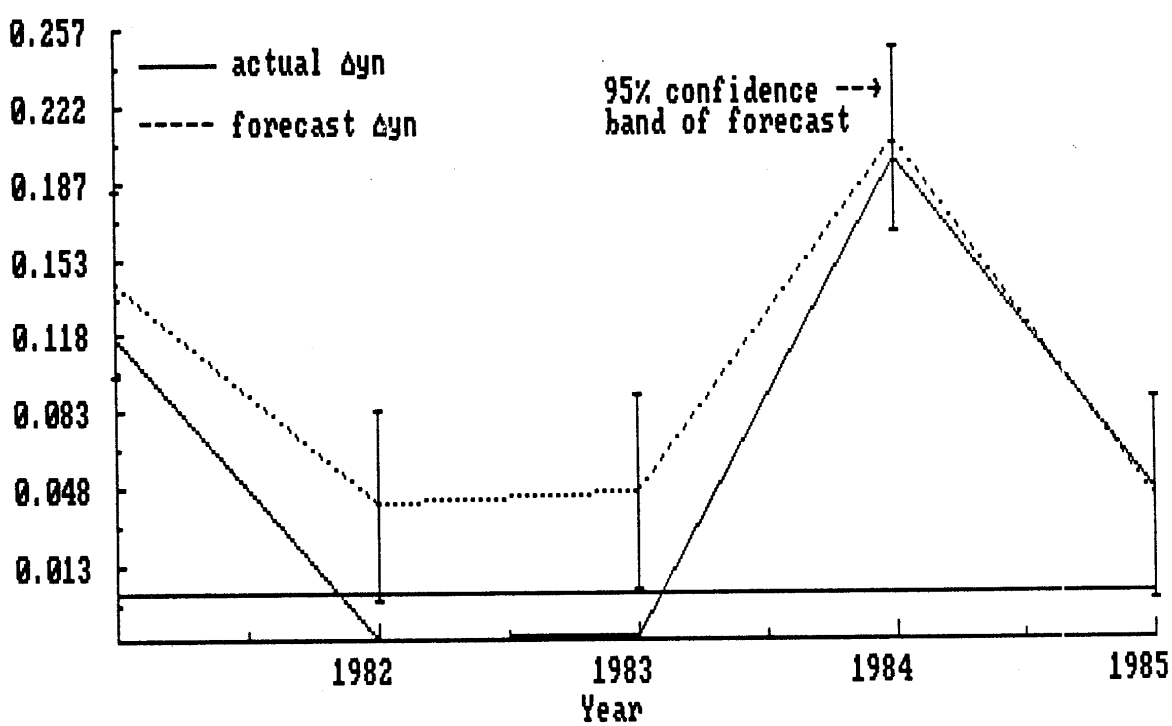


Figure 14. Equation (5.16) estimated over 1970-80: actual and forecast values of Δyn for 1981-85, and 95% confidence bands for the forecasts.



in (5.16) are non-constant: for (5.16) estimated over 1970-80, the one-period Chow statistics for 1982 and 1983 (η_{1_a} and η_{1_b}) are highly significant. Figure 14 plots actual and forecast values with 95% confidence bands for the forecasts: 1982 and 1983 are significant outliers.

To summarize, the marginal processes for prices and income are non-constant over the sample period yet the empirical model (5.4) of expenditure conditional on observed prices and income has constant parameters.[26] Thus, the parameters in (5.4) (and so the parameters of interest) are invariant to the class of interventions which occurred in sample: prices and income are super exogenous for those parameters.[27]

Super exogeneity has further ramifications. Because (5.15) and (5.16) have <u>non</u>-constant parameters, constancy of (5.4) implies that a broad class of expectations-based empirical models for expenditure <u>cannot</u> have a constant parameterization. This result, developed by Hendry (1988), is the focus of Section 6 on encompassing.

6. <u>Encompassing, Alternative Models, and Alternative Interpretations</u>

Encompassing, one of the primary criteria listed in Table 1, is loosely speaking the ability of one model to account for the results of another model. Omitted variables bias in linear models is a good example: the less

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[26]Equation (5.4) is a nonlinear function of p_t and Δyn_t in (5.15) and (5.16). Those variables proved simpler to model than Δp_t and Δy^*_t, which appear directly in (5.4). That does not affect the discussion on exogeneity.

[27]Just what the parameters of interest <u>are</u> is fundamental to any econometric investigation. Unfortunately, the notion of parameters of interest is taken for granted (and perhaps lost) in the standard textbook approach to exogeneity. Economic theory determines in part what are parameters of interest, e.g., ϕ and $\{\theta_i\}$ for the model in Section 2. Also, the empirical statistical properties of the estimated coefficients play a role, e.g., statistically constant coefficients generally are of more economic interest than non-constant ones when using data from an economy in which structural change and changing data moments are typical. The reason for the latter is economic interest in underlying ("structural") parameters. Structural invariance is a principal link between those economic-theoretic and statistical concepts; cf. Lucas (1976) and Engle <u>et al</u>. (1983).

restrictive (and, by assumption, correct) model predicts that the restricted model has biased coefficient estimates given by the standard formula. Encompassing applies to non-nested and nonlinear hypotheses as well, and to virtually any characteristic of a model (e.g., forecasts, error variances, autocorrelation in residuals, parameter non-constancy); cf. Mizon and Richard (1986). However, we are unaware of any alternative data-coherent empirical models for Venezuelan consumers' expenditure on recent data; cf. Appendix C. That <u>does not</u> leave encompassing as a (temporarily) vacuous criterion. First, by construction, (5.4) necessarily encompasses any empirical model nested in the model on Table 2. Second, because variance dominance is a necessary condition for variance and parameter encompassing, any new model must have an error variance of under .9% to encompass (5.4). Finally, because (5.15) and (5.16) are not constant but (5.4) is, broad classes of empirical models cannot encompass (5.4), <u>even in principle</u>. The remainder of this section investigates that last implication for constant-parameter expectations-based empirical models in general, for Hall's variant of the life cycle-permanent income hypothesis in particular, and for vector autoregressions (VARs).

To clarify why expectations-based empirical models cannot encompass (5.4), again consider the linear conditional model (3.1). Suppose that z_t is a valid instrumental variable for (3.1) and (3.1) holds (i.e., β <u>is</u> the parameter of interest). Instrumental variable estimation of (3.1) using z_t is consistent for β:

$$E(y_t|z_t) = \beta' E(x_t|z_t) \quad . \tag{6.1}$$

Thus, the conditional model (3.1) has an expectations interpretation. The converse, that expectations models have a conditional interpretation, is <u>not</u> true when the process for x_t changes and "conditional" means conditional upon <u>observed</u> x_t. As shown in Section 5, if the expectations model (5.9) holds, then

$$E(y_t|x_t) = \bar{\beta}'x_t \qquad (6.2)$$

where $\bar{\beta} = \bar{\beta}(\beta^*, \rho)$: $\bar{\beta}$ varies as ρ varies, even when β^* is constant. Conversely, if empirically ρ varies but $\bar{\beta}$ does not, the data cannot sustain the expectations interpretation (5.9) with β^* constant. That adds a twist to the empirical demonstration of super exogeneity. Stated more generally (and non-technically):

<u>Proposition</u>. If an empirical model for y_t conditional upon x_t has constant parameters, and a marginal process for x_t involves z_t and has parameters which change over time, then the data are not compatible with <u>any</u> expectations model for y_t with a set of constant parameters and for which the expectation of x_t depends upon z_t. See Hendry (1988).[28]

One might object that the proposition is dependent upon the particular estimated marginal process for x_t. Surprisingly, the proposition holds for <u>all</u> marginal processes for x_t, so long as they <u>include</u> z_t. For instance, for $z_t = x_{t-1}$ and $\rho_0 \neq \rho_1$ as in (5.13a-b), the proposition holds for all marginal processes for x_t which include x_{t-1}. The proof, given in Hendry (1988), is both ingenious and remarkably simple. If the marginal process for x_t based on some variable(s) z_t changes but extending the information set to include w_t (say) makes the marginal process for x_t constant, then the process for w_t conditional on z_t must change. Hence the "joint" marginal process for $z_t^* = (z_t : w_t)$ is non-constant as well. By assumption,

$$E(y_t|z_t^*) = \beta' E(x_t|z_t^*) \quad , \qquad (6.3)$$

so the proposition with z_t applies with z_t re-interpreted as z_t^*.

Again, the AR(1) process for x_t is illuminating. Suppose $\rho_0 \neq \rho_1$ in (5.13) but the "true" process for x_t requires some additional variable w_t:

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[28] Similar results apply when <u>future</u> values of x appear in expectations, but the details have not been completely worked out.

In practice, we do not know for certain whether or not a parameter is constant, so the empirical statements in the text are qualified by the properties of the statistics for testing parameter constancy.

$$x_t = \rho^* x_{t-1} + \tau' w_t + v_t^* \qquad v_t^* \sim NI(0,\sigma_*^2) \quad t=1,\ldots T \qquad (6.4)$$

where $E(x_{t-1}v_t^{*\prime})=0$ and $E(w_t v_t^{*\prime})=0$ by construction. Either (ρ^*,τ) vary over time or they are constant. If they vary, the proposition clearly holds. If they are constant, the marginal process for w_t must have changed; otherwise, ρ_i could not have changed in (5.13), noting the formula for omitted variable bias when excluding w_t from (6.4) to obtain (5.13). That implies that the marginal process for $(x_t:w_t)$ must have changed, so the proposition holds.[29]

The relevance of the proposition for the empirical results in Section 5 is the following. No empirical expectations-based model for Venezuelan consumers' expenditure on non-durables and services with y^* and p endogenous can encompass (5.4) if it includes any of Δp_{t-1}^+, Δp_{t-1}, or $\{xp_t, xp_{t-1}, xp_{t-2}, yn_{t-1}, yn_{t-2}\}$ in the set of information for forming expectations about y_t^* and p_t. In particular, it could not account for why the coefficient estimates in (5.4) are constant. Given the institutional structure of the Venezuelan economy, at least one of those variables almost invariably is an important determinant of y_t^* or p_t: hence expectations models for consumers' expenditure in Venezuela are unlikely to be consistent with the data.[30] However, even though a given expectations model is likely to be significantly mis-specified, (5.4) may be unable to encompass it: if so, that would indicate scope for improving (5.4) and might suggest how to do so. Such an encompassing test requires explicit specification of an expectations model and, until one exists or until other evidence is brought to light, (5.4) appears to be congruent with the data (see Hendry and

[29] Similar arguments apply for changes in equation error variances.

[30] This raises the question of why agents would use (condition upon) relatively simple functions of the data (including those of the forcing variables) rather than on the process of the forcing variables. One reason may be that, in a world with frequently and rapidly changing regimes of the forcing variables, data functions may provide relatively efficient and reliable predictions (or measures of) the underlying latent variables without the informational costs attendant to modeling the process of the forcing variables and the changes of that process.

Richard (1982, 1983)). We do <u>not</u> conclude that (5.4) is the "true" structural relationship, but its properties indicate its usefulness for inference and policy analysis, and it can serve as the basis for a progressive research strategy in which newer models encompass previous ones and account for additional as-yet-unexplained features of the data.

The life cycle-permanent income hypothesis is an important expectations-based model of consumers' behavior. Hall (1978) derives some statistical implications of that hypothesis, among which are: only the first lag of expenditure helps predict current expenditure and, conditional on that lag of expenditure, lagged disposable income offers no more predictive power, i.e., income does not Granger-cause expenditure. Estimating an AR(2) process for c_t yields:

$$\hat{c}_t = \underset{\substack{(.27)\\ [.25]}}{1.16}c_{t-1} - \underset{\substack{(.25)\\ [.25]}}{.26}c_{t-2} + \underset{\substack{(.54)\\ [.51]}}{.78} \tag{6.5}$$

T = 1970-1985 $R^2 = .9313$ $\hat{\sigma} = 4.498\%$ dw = 1.82 ADF(1) = -1.42 .

Although the t-statistic for c_{t-2} is -1.03 and insignificant by standard measures, two other features of (6.5) may invalidate that inference: the one-step Chow statistics for 1979 and 1984 of the AR(1) process for c_t indicate parameter non-constancy, and the Dickey-Fuller statistic for a unit root in (6.5) is "insignificant". Even ignoring these issues, deleting c_{t-2} and adding two lags of income gives:

$$\hat{c}_t = \underset{\substack{(.13)\\ [.15]}}{.88}c_{t-1} + \underset{\substack{(.10)\\ [.05]}}{.32}y^*_{t-1} - \underset{\substack{(.16)\\ [.12]}}{.23}y^*_{t-2} + \underset{\substack{(.47)\\ [.39]}}{.20} \tag{6.6}$$

T = 1970-1985 $R^2 = .9619$ $\hat{\sigma} = 3.485\%$ dw = 2.39 $\eta_3(2,12) = 5.7$

where η_3 tests for the significance of (y^*_{t-1}, y^*_{t-2}) and has a 95% nominal critical value of 3.89. Lagged income is highly significant, rejecting the second implication of the life cycle-permanent income hypothesis. That may not be surprising for Venezuela, given its income distribution, financial institutions, and the importance of liquidity in (5.4). However, the life cycle-permanent income hypothesis is a widely espoused theory of consumer behavior, so testing its implications bears some merit.[31] At a more general level, the results above on feedback versus feedforward models apply to (and reject) the life cycle-permanent income hypothesis because the latter is an expectations model and lagged inflation and petroleum revenues are not excluded from the information set for forming expectations in that model.

VARs are another common approach to modeling economic time series. Forecasting ability is a primary claimed advantage of VARs over other classes of models: even the proponents of VARs admit that the parameters estimated lack specific economic interpretation because they are reduced-form; see Sims (1987). For economies with structural change, VARs may no longer forecast well. For instance, the bivariate vector autoregressive representation of $(y_t:x_t)$ from (3.1) and (5.13) is:

$$y_t = \pi_i x_{t-1} + (u_t+\beta v_t) \tag{6.7a}$$

$$x_t = \rho_i x_{t-1} + v_t \tag{6.7b}$$

where $\pi_i=\rho_i\beta$. As ρ_i changes, so will the VAR coefficients $(\pi_i:\rho_i)$, and the entire VAR will manifest predictive failure; cf. Hendry (1979). Nonetheless, it may be difficult to detect parameter non-constancy because the error variance in (6.7a) is larger than that in (3.1) and because VARs tend to be heavily over-parameterized.[32]

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[31] Finite sample properties of statistics for testing "excess sensitivity" (e.g., η_3) have been widely discussed; cf. Banerjee and Dolado (1987) and Stock and West (1988). In (6.6), lagged income remains significant if a unit coefficient is imposed on lagged expenditure and/or if expenditure enters as a lagged rate of change only. Contrasting with Mankiw and Shapiro's (1985) Monte Carlo results, lagged income is insignificant if income appears at only one lag and a unit coefficient is imposed on lagged expenditure. Inclusion of D_{01} does not affect these results.

Parameter non-constancy (likely for 1981 and 1985) may compromise "classical" inferential procedures such as those which Hall uses. See also Davidson and Hendry (1981).

[32] Equations (5.4), (5.15), and (5.16) jointly imply that a VAR representation of expenditure such as (6.6) should have non-constant coefficients. Even though the residual variance in (6.6) is over <u>fourteen</u> times that in (5.4), the 1-step Chow statistics for (6.6) point to predictive failure in 1981 and 1985.

Factorization of a VAR into conditional and marginal models may mean more straightforward economic modeling of the time series. If the factorization implies super exogeneity, the conditional density may be amenable to a relatively parsimonious specification, as in (5.4). When the factorization corresponds to agents conditioning on outcomes from a marginal process, the marginal model (as well as the conditional one) may have an economic interpretation, thereby permitting easier modeling of the structural break(s) in the marginal model.[33] The approach in Section 5 is similar to that with VARs in that both allow a general dynamic specification initially. It differs from the latter in its emphasis on model evaluation and design, economically interpretable parameters of interest, and (relatedly) parsimony of the final specification and parameter constancy.

Having found (5.4) to be a tentatively adequate and economically interpretable characterization of the data, and having set aside a large class of alternative models, we turn to policy implications.

7. Policy Implications

Statistically valid policy analysis of an empirical model requires super exogeneity, i.e., that the parameters of interest β are invariant to a given class of interventions (changes) in the process for x_t. Section 5 demonstrates the super exogeneity of income and prices for the parameters in (5.4) with respect to the sorts of structural breaks and policy rule changes which occurred in sample. Before considering the policy implications of (5.4), some discussion of the Lucas critique is required.

The Lucas critique has been confirmed often in the empirical literature. It is less well-known that the Lucas critique can be refuted: Hendry (1988) implies refutability. In particular, the empirical model

[33] Lubrano, Pierse, and Richard (1986) illustrate this well for the determination of money and interest rates in the United Kingdom, where the introduction of Competition and Credit Control in 1971 provided the break.

(5.4) refutes the Lucas critique for the class of regime changes and policy rule switches which occurred in sample. Obviously, but non-trivially, the data are <u>un</u>informative for classes of interventions which did not occur. For instance, (5.4) might not have constant parameters if the personal income tax rate increased to 100%. Refutation of the Lucas critique is tempered because economic agents can modify their own behavior.

Given these caveats, (5.4) has the following policy implications. Foremost, government policy <u>does</u> have the potential to influence the path of real variables: that follows because (5.4) has a constant parameterization in spite of structural change within sample. In so far as government actions affect income, liquidity, and inflation, the government can (and will) affect consumers' expenditure. The dependence of income on petroleum exports (and on imports of manufactured goods) is clear: it had dramatic repercussions for the economy, both positively and negatively (e.g., 1974 and 1982-83). Thus, diversification of exports and expansion of the domestic manufacturing industry could reduce the potential volatility of consumers' expenditure. Price (or wage and price) policies and policies influencing real income growth and liquidity could affect both current and equilibrium expenditure through $(\Delta p+\Delta^2 p)_t$, $(\Delta_2 y_t^*/2)$, y_t^* itself (which is $y_t-\Delta p_t \cdot \ell_{t-1}$), and the error-correction term (which depends upon lagged y^* and ℓ).[34] To the extent that agents are liquidity-<u>constrained</u>, expansionary government fiscal policy may relax that constraint if trade-offs in utility exist between consumption of private and public goods; cf. Rossi (1988). A more detailed discussion of the policy implications of (5.4) requires specifying the policy at hand and analyzing the mechanism by which it influences prices, income, liquidity, and other economic variables.

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[34] However, price polices might not, to the extent that they implied rationing.

8. Concluding Remarks

This paper starts with a simple optimizing model in which economic agents aim to maintain constant expenditure-income and liquidity-income ratios but face adjustment costs. That theoretical model provides the basis for constructing a parsimonious, statistically adequate econometric model of consumers' expenditure on non-durables and services in Venezuela. Disposable income, liquidity, and inflation help determine expenditure, with all variables doing so through several economically sensible channels. To the model's benefit, its parameters are constant in spite of structural change in the economy: that evidence also excludes a large class of expectations and VAR models from being empirically viable. Rather than being the econometrician's bane, structural change proves invaluable for sorting through alternative models. The structure of the consumption function has numerous implications for economic forecasting and policy.

At another level, this paper illustrates an econometric methodology which appears promising for evaluating empirical models, designing them to have desirable statistical and economic properties, and improving the chances of reliable inference. The methodology applies equally to other <u>classes</u> of models: selecting a conditional model was an empirical issue, not one of principle. Further, the methodology provides means for resolving conflicts between competing empirical explanations of the same phenomena, thereby focusing empirical analysis into a progressive research strategy.

Appendix A. Data Definitions

VARIABLE	DEFINITION	SOURCE
ABT	Gasto de consumo final de los hogares en el mercado interno: alimentos, bebidas y tabaco [Consumers' expenditure on food, drink and tobacco]	A.C.N. III-5
C	Real per capita consumers' expenditure on non-durables and services (= [ABT+OT+SD+S]/[P•N]) (1968 Bolívares per capita) [Gasto real per capita de consumo en bienes no durables y servicios]	-
D_{01}	A dummy variable for mis-measurement on consumers' expenditure: +1 for 1970, -1 for 1971, 0 otherwise	See App. B
L	Liquidez monetaria (M_2) = monedas + billetes + depósitos a la vista + depósitos de ahorro + depósitos a plazo (end-of-year) [Monetary aggregate M_2 = coins + bills + sight deposits (checking) + savings deposits + time deposits (CDs)]	B.M. III.2.1
N	Population (mid-year estimates in millions) [Población]	I.F.S. 99Z
OT	Gasto de consumo final de los hogares en el mercado interno: otros bienes no durables [Consumers' expenditure on other non-durables]	A.C.N. III-5
P	Indice de precios al consumidor para el area metropolitana de Caracas; indice general (diciembre, 1968=1.00) [Consumer price index for the Caracas metropolitan area, December values]	B.M. III.4.6
P^+	Consumer price index for the United States (1980=1.00) [Indice de precios al consumidor para los EE.UU.]	I.F.S. 64
S	Gasto de consumo final de los hogares en el mercado interno: servicios [Consumers' expenditure on services]	A.C.N. III-5
SD	Gasto de consumo final de los hogares en el mercado interno: bienes semidurables [Consumers' expenditure on "semi-durables"]	A.C.N. III-5
XP	Petroleum exports [Exportaciones de petróleo]	I.F.S. 70A
Y	Real per capita national disposable income (= YN/[P•N]) [Ingreso nacional disponible, real per capita]	-
Y^*	Real per capita national disposable income, adjusted for the inflation tax on liquid assets [Ingreso nacional disponible, real per capita]	See text
YN	Ingreso nacional disponible [National disposable income]	A.C.N. I-1.3

Sources. The data sources are: A.C.N., _Anuario de Cuentas Nacionales_ (Annual Report of the National Accounts); B.M., _Boletín Mensual_ (Monthly Bulletin); and I.F.S., _International Financial Statistics 1987 Yearbook_. The first two are publications of the Banco Central de Venezuela (Caracas, Venezuela) and the last is published by the International Monetary Fund (Washington, D.C.). Data from the A.C.N. are from the 1982 volume, with extensions and revisions to the data from subsequent issues; data from the B.M. are from various issues. The number below a source indicates the relevant table (for A.C.N. and B.M.) or line in the corresponding country table (for I.F.S.). The definition as it appears in the source is given first. An English (or Spanish) translation appears in brackets.

Units. Unless otherwise noted, all published data are annual, and nominal in millions of Bolívares.

Adjustments. Population data before 1975 is rescaled to account for a re-definition in 1975; cf. _International Financial Statistics 1984 Yearbook_.

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Appendix B. Data Measurement

This appendix records several _caveats_ about the principal series analyzed in the text.

Consumers' expenditure (C) is final expenditure in the domestic market by households and non-profit institutions. It includes purchases in the domestic market by non-residents and excludes those by residents in the foreign market. In 1974, the former were at most 2% and the latter at most 8% of final expenditure in the domestic market.

There appear to be measurement errors in C for 1970 and 1971: Figures B.1, B.2, and B.3 graph the logs of nominal, real, and real per capita expenditure for 1968-1973. We have found no economic or institutional

Figure B.1. The log of nominal consumers' expenditure (c+p+n) for 1968-73.

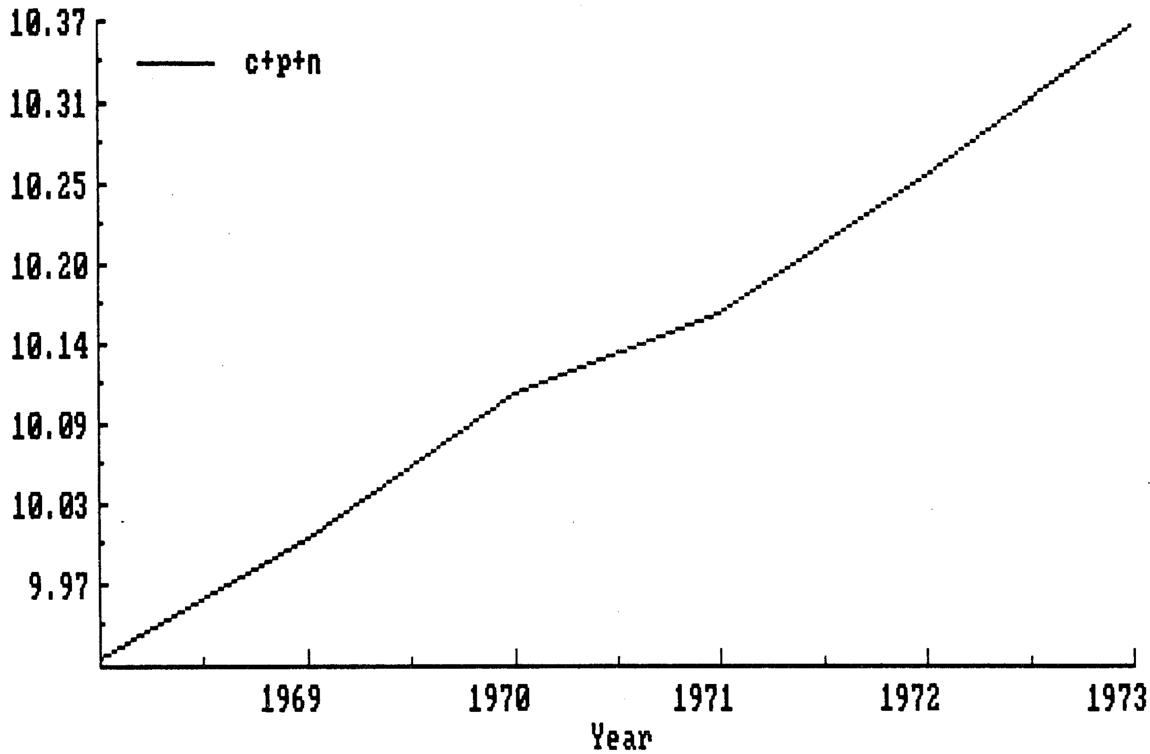


Figure B.2. The log of real consumers' expenditure (c+n) for 1968-73.

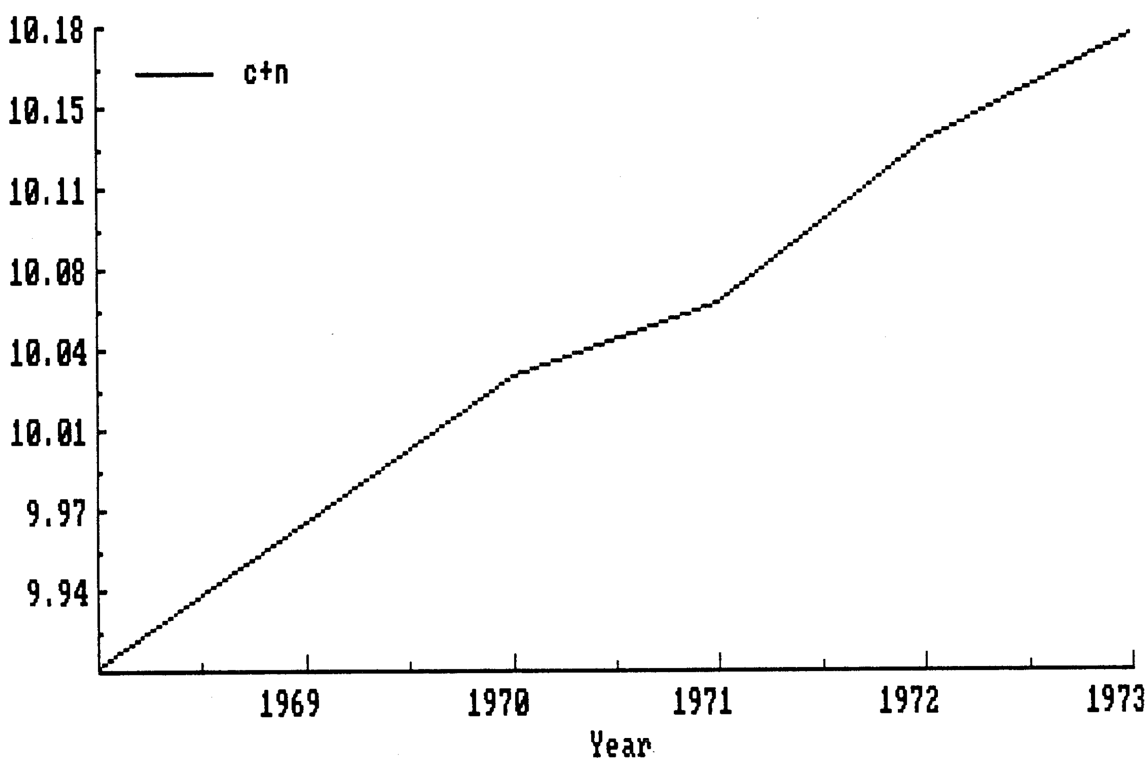


Figure B.3. The log of real per capita consumers' expenditure (c) for 1968-73.

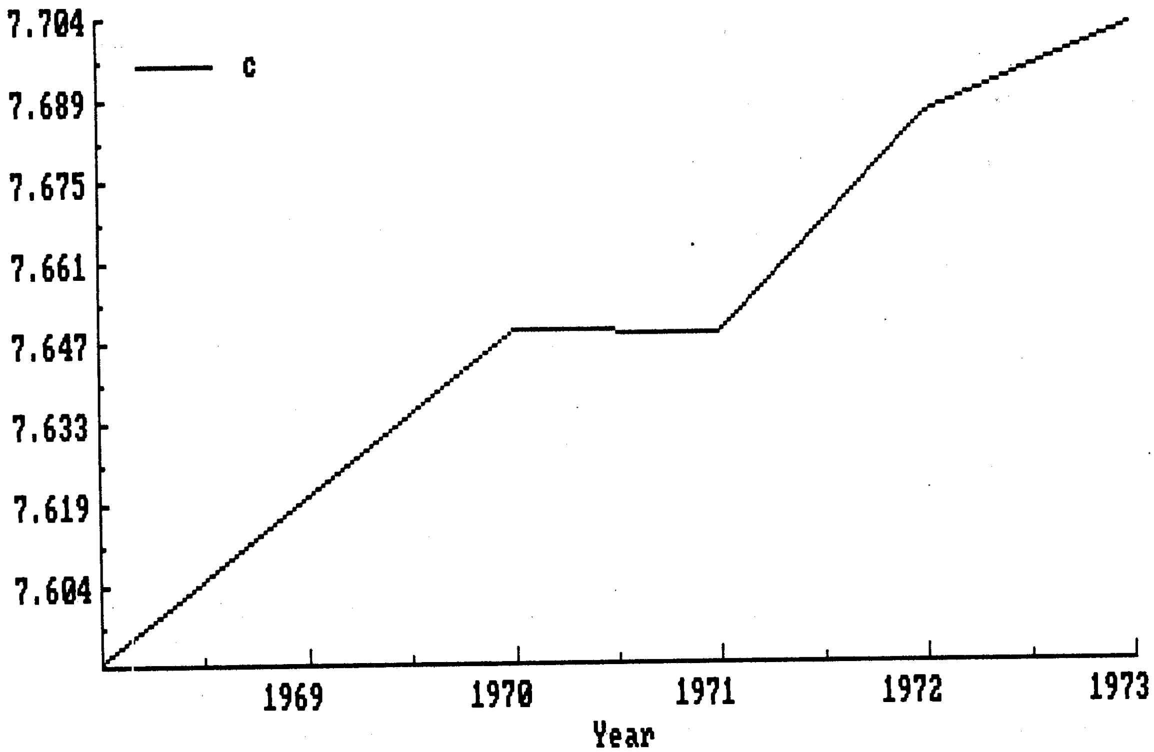
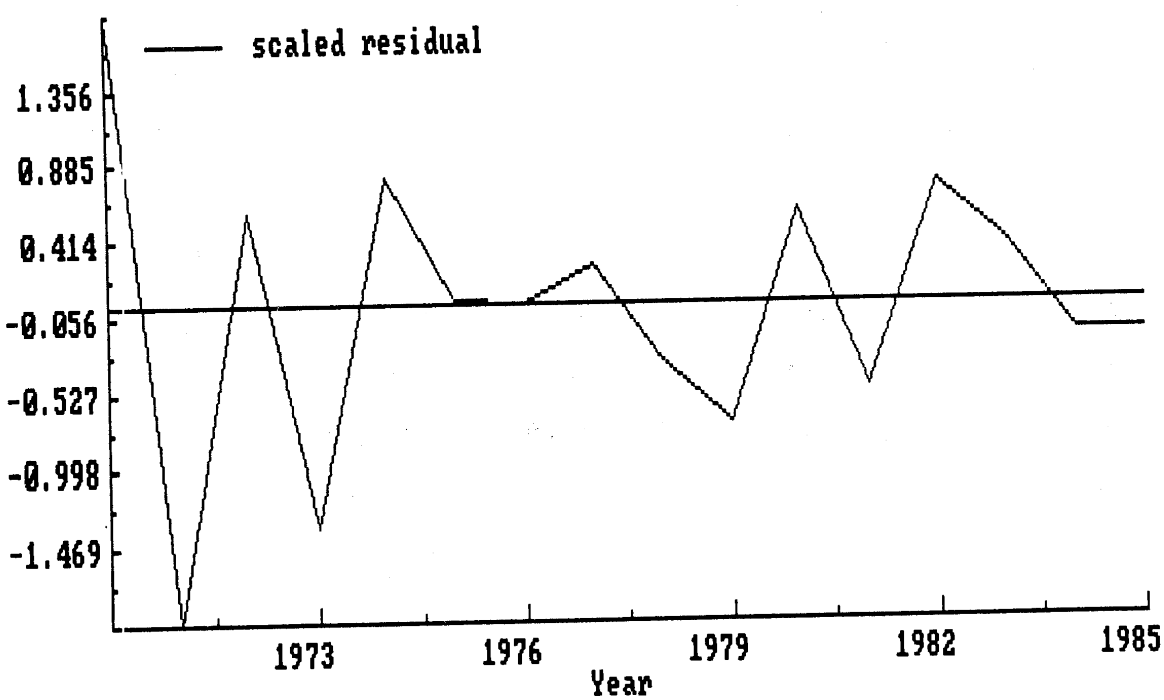


Figure B.4. Equation (B.1): scaled residuals from the error-correction equation (5.4) without D_{01}.



explanations of the unusual movements in 1970-71. Estimation of (5.4) without D_{01} results in virtually identical estimates of the remaining coefficients (see (B.1) below), but larger estimated standard errors and $\hat{\sigma}$. That is consistent with measurement errors on expenditure which are nearly orthogonal to the RHS variables in (B.1).

$$\begin{aligned}
\hat{\Delta c}_t = \; & \underset{\substack{(.046)\\ [.032]}}{.451}(\Delta_2 y^*_t/2) - \underset{\substack{(.039)\\ [.025]}}{.262}(\Delta p+\Delta^2 p)_t - \underset{\substack{(.028)\\ [.041]}}{.139}[(c-y^*)-.5(\ell-y^*)]_{t-1} \\
& + \underset{\substack{(.006)\\ [.010]}}{.019} \qquad\qquad (B.1)
\end{aligned}$$

$T = 1970\text{-}1985 \quad R^2 = .9317 \quad \hat{\sigma} = 1.3816\% \quad dw = 2.92 \quad \xi_2(3) = 17.1$

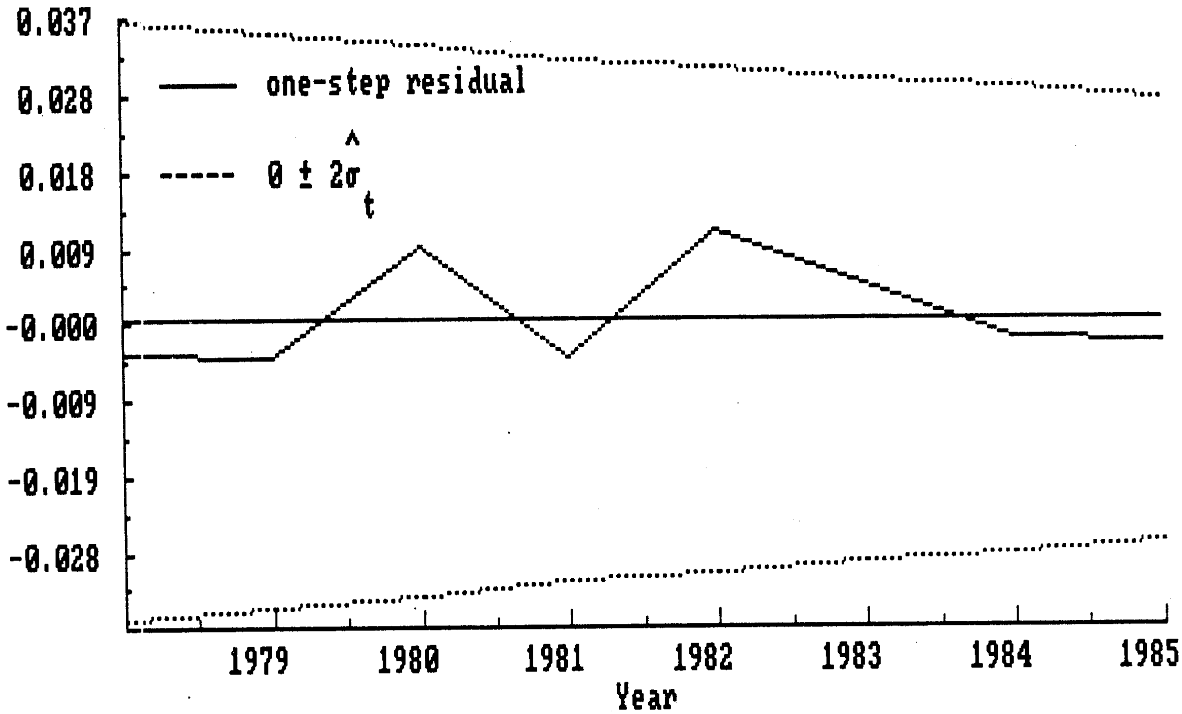
$\eta_4(1,10) = 3.96 \quad \eta_5(6,5) = 4.41 \quad \eta_6(2,10) = .4 \quad \xi_7(2) = .1$

Figure B.4 graphs the scaled residuals from (B.1): those for 1970 and 1971 are opposite in sign and nearly equal in magnitude, and all other residuals are considerably smaller. Figure B.5 plots the <u>one-step</u> residuals with calculated equation standard error bands. Those bands <u>narrow</u> over time, indicative of the relatively large outliers early on in the sample.

One explanation of these anomalies is a minor re-definition of the expenditure series in 1971. Another is an over-estimate of expenditure in 1970 and a compensating under-estimate in 1971. Both are plausible and, until we have additional evidence, we adopt the latter model, capturing the effect on C with the dummy D_{01}.

<u>National disposable income</u> (Y) is not the best measure of income conceivable; personal disposable income would be better, but it is not available. In particular, national disposable income includes profits of the petroleum industry, but those profits should not affect consumers' expenditure directly. Those profits were unusually high in 1973 and low in 1982-3: that may be responsible for (5.4) over-predicting in 1973 and under-predicting in 1982-3; cf. Figure 6.

Figure B.5. Equation (B.1): one-step residuals and the corresponding calculated standard errors.



The consumer price index (P) is for the Caracas metropolitan area only, so it does not reflect regional variations in prices. The index is for all expenditure, both durable and non-durable. Although the index may be well-suited for deflating expenditure of non-durables and services for our econometric analysis, the price of durable relative to non-durable goods also could be important, especially if $(\Delta p+\Delta^2 p)_t$ in (5.4) is interpreted as in Deaton (1977). The index is sensitive to the particular basket of goods used in its calculation and to the presence of price controls (which are common in Venezuela and have existed in varying degrees).

Liquid assets (L) are measured by the monetary aggregate M_2 and include holdings by both the personal and commercial sectors (but not the government and excluding financial institutions [las Sociedades Financieras]). They exclude holdings in Savings and Loans Associations ([Cédulas Hipotecarias]; very liquid) and holdings abroad. The latter increased dramatically over the sample because of capital flight. While accounting for capital flight may be important, it is not immediately obvious how (e.g.) Venezuelan assets in dollar accounts in the U.S. affect expenditures in Venezuela.

Appendix C. Encompassing, and Previous Empirical Models of Consumers' Expenditure in Venezuela

In contrast to most industrialized countries, little empirical research exists on consumer behavior in Venezuela. Belandria (1971) analyzes the demand for various goods with cross-section data from three cities (Barinas, Mérida, and San Cristóbal). He estimates three equations using income and either total expenditure or prices, and finds those variables significant. He rejects the hypotheses that corresponding coefficients are equal for the three cities. López (1972) estimates a subset of Zellner's (1957) equations and tests for habit formation, inertia, and the importance of liquid assets and of past levels of expenditure. López finds that (i) including lagged expenditure as an explanatory variable reduces the estimated marginal propensity to consume and increases the latter's standard error, (ii) the coefficient of liquid assets is very high (much more so than Zellner's), and (iii) the marginal propensity to consume typically is negative when estimated by indirect least squares. Palma (1976) builds a model of the Venezuelan economy which explains inter alia total consumers' expenditure, expenditure on food, beverages and tobacco, and expenditure on other non-durables and services for 1950-69. He finds that both liquidity and (a constructed series of) personal disposable income are important as explanatory variables. Musgrove (197*) examines cross-section data from a household budget survey to finds that total expenditure and family structure (number of members, age and sex) have a significant effect on expenditure. Finally, Villoria (1983) estimates several equations for total consumption in terms of current and lagged national income, lagged consumption and liquidity.

These studies use data over periods different from ours and gathered on different accounting principles.[35] Even so, Davidson et al. (1978) were successful in resolving similar issues in their analysis of different models of consumers' expenditure in the United Kingdom. Palma's (1976) data are perhaps closest to ours so, at a minimum, we plan to re-estimate (5.4) with his data and test if either his model or (5.4) can encompass the other.

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[35]In 1970, the Central Bank of Venezuela (which produces much of the Venezuelan economic data) adopted the then new United Nations system of national accounts; it has published data from 1968 according to those methods. Of the authors mentioned above, only Villoria uses time series under that system, and he considers total expenditure rather than expenditure on non-durables and services. Given the marked differences in the time-series behavior of expenditure on durables in comparison to that of non-durables and services, we chose to analyze the more disaggregated series. Encompassing his results is feasible with a model of durable expenditure in addition to one for non-durables and services.

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