VARIABILITY OF DURABLE AND NONDURABLE CONSUMPTION: EVIDENCE FOR SIX O.E.C.D. COUNTRIES

Jordi Galí*

Abstract—We estimate consumption variability ratios for both durables and nondurables consumption, using data for six OECD countries Our methodology, which relies on a long-run restriction implied by the consumer's intertemporal budget constraint, overcomes many of the problems inherent to previous approaches. Some important departures from the permanent income model emerge: (i) nondurables consumption shows mild excess smoothness in the United States and Italy, and mild excess volatility in Japan and France, and (ii) durables consumption shows extreme excess smoothness in all countries Alternative factors capable of generating the differences in volatility across types of goods are discussed.

I. Introduction

MODERN formulations of the permanent income model, as developed in Hall (1978) and Flavin (1981), provide an essential reference framework for understanding the dynamic behavior of consumption and savings. Because of the important role played by those variables in business cycle models, a large amount of work has been devoted to developing tests of the permanent income hypothesis (PIH), and to characterizing and interpreting the rejections of that hypothesis detected in the data.

Much of the recent work has focused on two departures from the permanent income model: (i) consumption changes are predictable (Flavin (1981)), thus violating the martingale property implied by the PIH, and (ii) consumption appears to be smoother than is implied by the permanent income model (Deaton (1987)). One can interpret many recent papers in this area as attempts to explain both puzzles by introducing features like liquidity constraints (Campbell and Mankiw (1989)), unobservable components in income (Quah (1990)), unrestricted information sets (Campbell and Deaton (1989), West (1988)), and finite horizons (Clarida (1988), Galí (1990)), among others.

In the present paper we revisit and extend the evidence on aggregate consumption's relative variability, i.e., the extent to which aggregate consumption is smoother or more volatile than is justifiable on the basis of news about current and future income.

Our approach builds on Galí (1991). The econometric strategy developed in that paper allows one to identify and estimate measures of permanent income variability which rely on relatively weak assumptions and thus avoid some of the shortcomings of previous approaches. Briefly, Galí (1991) shows that, as long as the intertemporal budget constraint is met, the autocorrelogram of consumption changes provides all the information required to make inferences on the variability of unobservable permanent income innovations. The results in Galí (1991), based on postwar U.S. data, imply that the variability in nondurables and services consumption is about 70% of that predicted by the PIH model. Simple departures from the basic version of that model —e.g., habit formation or liquidity constraints— were shown to be capable of explaining the observed excess smoothness.

Here we extend the analysis in Galí (1991) in two main directions. First, we show how the basic framework can be modified to simultaneously encompass consumption of both *durables* and *nondurables*, where the latter will henceforth include both nondurable goods and services. Under some assumptions, we can construct two statistics that measure the extent of excess smoothness or excess volatility in *nondurables* and *durables* consumption independently. Second, we apply our approach to consumption data for the United States, as well as a number of O.E.C.D. countries for which appropriate data are available. That cross-country evidence is used, following Jappelli and Pagano (1989) and others, to assess the plausibility of the liquidity constraints hypothesis.

Received for publication June 7, 1991. Revision accepted for publication June 19, 1992.

* Columbia University

A previous version of this paper was circulated under the title "International Evidence on Consumption Variability" I am grateful to participants at the CEPR International Macroeconomics Programme Meeting (El Escorial, 1991), the discussant Samuel Bentolila and two referees for useful comments and suggestions. I also thank Joon-Ho Hahm for valuable research assistance. I am solely responsible for any remaining errors and misinterpretations.

II. A Theoretical Framework for the Time Series Analysis of Durables and Nondurables Consumption

A. A Benchmark Permanent Income Model

We start by developing a simple version of the permanent income model which integrates both durable and nondurable goods consumption.

The representative consumer solves the following optimization problem:

$$\max E_0 \sum_{t=0}^{\infty} (1+\rho)^{-t}(U_t(cn_t) + F_t(k_t))$$

subject to

$$w_{t+1} = w_t(1+r) + y_t - (cn_t + cd_t) \tag{1}$$

$$k_t = k_{t-1}(1-\delta) + cd_t \tag{2}$$

$$cn_t, k_t \geq 0;\ \lim_{T\to\infty} E_0(1+r)^{-T} w_T = 0 \tag{3}$$

where cn is consumption of nondurables, cd is expenditures on durable goods, k is the stock of durable goods, w is nonhuman wealth (excluding k), δ is the depreciation rate, and y is labor income. The rate of return on nonhuman wealth is assumed to be constant, and is denoted by r. Notice that our specification of preferences assumes separability between cn and k, while allowing for time-dependent utility functionals. Units of capital are normalized so that their price (by assumption constant) equals the price of the nondurable good, which in turn is normalized to unity.

The solution to the problem above satisfies the Euler equation

$$U'_{t-1}(cn_{t-1}) = ((1+r)/(1+\rho))E_{t-1}U'_t(cn_t) \tag{4}$$

and an intratemporal efficiency condition

$$F'_t(k_t) = \tau U'_t(cn_t) \tag{5}$$

relating the marginal rate of substitution between k and cn to the implicit rental cost $\tau \equiv (r+\delta)/(1+r)$.[1]

When both (4) and (5) hold, it follows that

$$F'_{t-1}(k_{t-1}) = ((1+r)/(1+\rho))E_{t-1}F'_t(k_t). \tag{6}$$

[1] Assuming an interior solution, conditions (4) and (5), combined with the transversality conditions $\lim_{T\to\infty} E_0(1+\rho)^{-T}U'_T(cn_T)W_T = 0$ and $\lim_{T\to\infty} E_0(1+\rho)^{-T}U'_T(cn_T)k_T = 0$, are necessary and sufficient

We specialize our model by assuming $r = \rho$, and the following specification for preferences:

$$U_t(cn_t) \equiv -0.5(\chi_t - cn_t)^2; \qquad \chi_t \equiv \chi_0 + \alpha_n t$$

$$F_t(k_t) \equiv -0.5\nu(\Omega_t - \theta k_t)^2; \quad \Omega_t \equiv \Omega_0 + \alpha_k t$$

where θ is the amount of "services" per period provided by one unit of durable goods, and χ_t and Ω_t can be interpreted as consumption bliss levels. Under the previous assumptions, (4) and (6) can be rewritten as

$$\Delta cn_t = \alpha_n + \xi_{nt} \tag{7}$$

$$\Delta cd_t = \alpha_d + (1-(1-\delta)L)\xi_{dt} \tag{8}$$

where $\alpha_d \equiv (\delta\alpha_k/\theta)$. By construction, $\xi_{nt} \equiv cn_t - E_{t-1}cn_t$ and $\xi_{dt} \equiv cd_t - E_{t-1}cd_t$ are the innovations in nondurables and durables consumption, respectively. (7) is just the well-known martingale result originally derived in Hall (1978), whereas (8) is its durable goods version, introduced in Mankiw (1982).

We can rewrite (5), after some manipulation, as follows:

$$cd_t = (\delta\gamma/\theta)cn_t + u_t \tag{9}$$

where $\gamma \equiv (\tau/\nu\theta)$, and

$$u_t \equiv (\delta/\theta)(\Omega_0 - \gamma\chi_0) + [(1-\delta)\alpha_d/\delta] + [\alpha_d - (\delta\gamma/\theta)\alpha_n]t + (\gamma/\theta)(1-\delta)\xi_{nt}.$$

Since u is a stationary process (around a deterministic trend) it follows that cn and cd are "stochastically cointegrated" with a cointegrating vector $[1, -(\delta\gamma/\theta)]$.

The previous cointegration result, together with the intertemporal budget constraint, allows us to pin down the relationship between innovations ξ_n and ξ_d, and their "fundamental" determinant, i.e., news about future labor income. At this point it is useful to introduce the notion of permanent income yp, as defined in Flavin (1981). Letting $\lambda \equiv 1/(1+r)$,

$$yp_t \equiv r\left(W_t + \lambda \sum_{j=0}^{\infty} \lambda^j E_t y_{t+j}\right).$$

The innovation in permanent income, $\xi_t \equiv yp_t - E_{t-1}yp_t$, is thus given by

$$\xi_t = r\lambda \sum_{j=0}^{\infty} \lambda^j (E_t - E_{t-1})y_{t+j}.$$

Such news about current and future income is the only source of consumption fluctuations in the model above.

The dynamic budget constraint (1) and the transversality condition in (3) can be used to derive the standard intertemporal budget constraint. Applying the law of iterated expectation to the latter, we get

$$r\lambda \sum_{j=0}^{\infty} \lambda^j (E_t - E_{t-1})(cn_{t+j} + cd_{t+j}) = \xi_t, \tag{10}$$

which has a simple interpretation: if the intertemporal budget constraint is satisfied, the present discounted value of revisions in expected future consumption must match the innovation in total wealth ξ/r.

We postulate a simple linear relationship between innovations in *cn* and *cd* and contemporaneous innovations in permanent income:

$$\xi_{nt} = \beta_n \xi_t; \qquad \xi_{dt} = (1/\delta)\beta_d \xi_t$$

where β_n and β_d are constants to be determined.[2] The permanent effect on *cn* of a one unit innovation in permanent income is thus β_n. The corresponding effect on *cd* is given by β_d. The cointegration result above thus implies $\beta_d = (\delta\gamma/\theta)\beta_n$.

An additional restriction on the β's, resulting from the need to satisfy the intertemporal budget constraint, follows from (7), (8) and (10):

$$\beta_n + (\tau/\delta)\beta_d = 1.$$

β_n and β_d can now be solved for in terms of the exogenous parameters of the model:

$$\beta_n = [\theta/(\theta + \tau\gamma)]; \qquad \beta_d = [\delta\gamma/(\theta + \tau\gamma)].$$

Letting $\sigma_*(\Delta cn)$ and $\sigma_*(\Delta cd)$ denote the standard deviations of Δcn and Δcd implied by the permanent income model above, we have

$$\sigma_*(\Delta cn) = [\theta/(\theta + \tau\gamma)]\sigma(\xi) \tag{11}$$

$$\sigma_*(\Delta cd) = \sqrt{1 + (1-\delta)^2} \times [\gamma/(\theta + \tau\gamma)]\sigma(\xi), \tag{12}$$

which pin down $\sigma_*(\Delta cn)$ and $\sigma_*(\Delta cd)$ as functions of the exogenous parameters r, δ, θ, ν, and $\sigma(\xi)$.

Our goal is to compare $\sigma_*(\Delta cn)$ and $\sigma_*(\Delta cn)$ with the corresponding empirical standard deviations of nondurables and durables consumption, which we denote by $\sigma(\Delta cn)$ and $\sigma(\Delta cd)$. With that purpose in mind we introduce the following statistics:

$$\varphi_n \equiv \sigma(\Delta cn)/\sigma_*(\Delta cn)$$

$$\varphi_d \equiv \sigma(\Delta cd)/\sigma_*(\Delta cd).$$

Values of φ less than one are naturally interpreted as evidence of *excess smoothness*, whereas values greater than one correspond to *excess variability*. Unfortunately, estimation of φ-type statistics faces a major difficulty: $\sigma_*(\Delta cn)$ and $\sigma_*(\Delta cd)$ are not observable. Even if we knew and/or were able to estimate β_n and β_d, there would still remain the question of how to identify and estimate $\sigma(\xi)$. This is especially problematic because neither permanent income *yp* nor its innovation ξ are observable, since they depend on (unobservable) expectations. That problem is often tackled in the literature by assuming that some transformation of labor income follows a univariate ARMA process.[3] That approach has two important shortcomings: (i) the estimates of φ-type variance ratios are extremely sensitive to the assumption made on the order of integration of labor income, i.e., on the particular transformation of that variable used, and (ii) the approach implicitly (and arbitrarily) assumes that consumers use only information contained in current and past values of labor income in order to predict future labor income.

In West (1988) and Campbell and Deaton (1989), a methodology is developed to test the PIH using estimates of φ-type ratios, without imposing any restriction on the consumer's information set. However, as stressed in Flavin (1988), the estimated variability ratios *are only correct under the PIH null*, so they should be used only for the purpose of testing that null.[4]

The approach taken in this paper, originally developed in Galí (1991) and extended here to handle durables consumption, does not require any restriction on the information set used by consumers to form their expectations about future income. Its main advantage relative to the West and Campbell-Deaton approaches lies in its *robustness to a wide range of departures from the*

[2] Premultiplication of β_d by $(1/\delta)$ is just a convenient normalization

[3] For a further discussion on this issue, with complete references, see Galí (1991)

[4] The outcome of those tests systematically rejects the PIH model for nondurables, thus implying that the estimates of φ obtained by both the West and the Campbell-Deaton procedures, *though sufficient to reject the PIH null, are no longer "admissible" estimates of the true* φ

PIH model. Additionally, that approach does not rely on any assumption on the stochastic properties of the labor income process.[5]

B. Departures from the Permanent Income Model and Identification of Variability Ratios

As a rather unrestricted alternative to the permanent income model specified above we postulate the following models for Δcn and Δcd:

$$\Delta cn_t = \alpha_n + \beta_n(L)\xi_t + \phi_n(L)\eta_t \qquad (13)$$

$$\Delta cd_t = \alpha_d + \beta_d(L)\xi_t + \phi_d(L)\nu_t \qquad (14)$$

where $\beta_n(L)$, $\beta_d(L)$, $\phi_n(L)$, and $\phi_d(L)$ are polynomial functions in the lag operator L.

Thus, two different sources of consumption movements are allowed for: shocks to permanent income (ξ) and preference shocks (η and ν). We assume that the latter are serially uncorrelated, and uncorrelated with innovations in permanent income at all leads and lags. The long-run effect of a permanent income innovation on cn is given by $\beta_n(1)$, whereas the corresponding effect on cd is equal to $\beta_d(1)$. Notice that (13) and (14) nest the permanent income model developed in the previous section, which corresponds to $\beta_n(L) = (\theta/(\theta + \tau\gamma))$, $\beta_d(L) = (\gamma/(\theta + \tau\gamma))\ (1 - (1 - \delta)L)$, and $\phi_n(L) = \phi_d(L) = 0$.

Even if the permanent income model does not hold, the intertemporal budget constraint and, consequently, equation (10) must be satisfied for all possible realizations of ξ, η, and ν. Accordingly, the following restrictions on the parameters of consumption processes (13) and (14) must hold:[6]

$$\beta_n(\lambda) + \beta_d(\lambda) = 1 \qquad (15)$$

$$\phi_n(\lambda) = \phi_d(\lambda) = 0. \qquad (16)$$

Assuming continuity of the polynomial functions $\beta_n(\lambda)$, $\beta_d(\lambda)$, $\phi_n(\lambda)$, and $\phi_d(\lambda)$, at $\lambda = 1$, the following limit result obtains:

$$\lim_{r\to 0} (\beta_n(1) + \beta_d(1)) = 1 \qquad (17)$$

$$\lim_{r\to 0} \phi_n(1) = \lim_{r\to 0} \phi_d(1) = 0 \qquad (18)$$

where the limit is taken for a sequence of economies whose associated interest rates converge to zero.

We need to make a further assumption: in the absence of preference shocks, efficiency condition (5) is assumed to hold at zero frequency. In other words, we assume that the eventual presence of short-run deviations from the PIH does not prevent consumers from equating the marginal rate of substitution between durables and nondurables to their relative price, in the long run, and absent preference shocks. Accordingly, the cointegration result in (9) will still hold, thus implying

$$\beta_d(1) = (\delta\gamma/\theta)\beta_n(1). \qquad (19)$$

Combining (17) and (19), and given that $\lim_{r\to 0} \tau - \delta = 0$, we get

$$\lim_{r\to 0} \{\beta_n(1) - [\theta/(\theta + \tau\gamma)]\} = 0 \qquad (20)$$

$$\lim_{r\to 0} \{\beta_d(1) - [\delta\gamma/(\theta + \tau\gamma)]\} = 0 \qquad (21)$$

which will play a central role in what follows.

The following lemma is the basis for much of the analysis below.

Lemma: Let $f_n(\omega)$ and $f_d(\omega)$ be the normalized spectra of Δcn and Δcd, respectively. Define $\mu(\delta) \equiv \delta/\sqrt{1 + (1 - \delta)^2}$. Under the consumption model given by (13) *and* (14) *the following limit results hold*:

$$\lim_{r\to 0} \left[\left\{1/\sqrt{2\pi f_n(0)}\right\} - \varphi_n\right] = 0 \qquad (22)$$

$$\lim_{r\to 0} \left[\left\{\mu(\delta)/\sqrt{2\pi f_d(0)}\right\} - \varphi_d\right] = 0. \qquad (23)$$

Proof: See appendix.

In practice, and given the continuity of our polynomial functions, $\{1/\sqrt{2\pi f_n(0)}\} \cong \varphi_n$ and $\{\mu(\delta)/\sqrt{2\pi f_d(0)}\} \cong \varphi_d$ will generally be good approximations for plausibly low interest rates.[7] The lemma above thus implies a strong connection between the consumption variability ratios φ_n and φ_d introduced in the previous section and the normalized spectrum of the durable and non-

[5] That has two advantages. First, we do not need to make any assumption on the order of integration of labor income Second, our results are robust to the presence of different components in labor income unobservable to the econometrician, thus overcoming the problem pointed out by Quah (1990)

[6] In (15) we are implicitly assuming that η_t and ν_t are not perfectly correlated Otherwise, i e., if $\eta_t = \pi\nu_t$, all t, the corresponding restriction would be $\phi_n(\lambda)\pi + \phi_d(\lambda) = 0$, but $\phi_n(\lambda)$ and $\phi_d(\lambda)$ would not be independently restricted.

[7] See Galí (1991) for some numerical examples illustrating the accuracy of the approximation.

durable consumption processes. Identification of $\sigma_*(\Delta cn)$ and $\sigma_*(\Delta cd)$—i.e., the measures of consumption variability under the PIH model—relies on three basic assumptions: (i) the intertemporal budget constraint is met with equality, and (ii) there is a linear stationary representation for consumption changes, in terms of a distributed lag of permanent income innovations and preference shocks, and (iii) the efficiency condition relating the MRS between durables and nondurables to their relative price is satisfied, at least in the long run. Assumptions (i) and (ii) played a central role in the identification of $\sigma(\xi)$ in Galí (1991) in a model with only nondurable goods. Assumption (iii) makes it possible to extend that analysis to the case of durables.

Given the results above, we can use any of the available consistent estimators of the normalized spectrum at frequency zero $f_n(0)$ to construct a consistent estimator for φ_n. A similar strategy can be followed in order to estimate φ_d, though in the latter case we must condition the estimate on a given value of δ, the depreciation rate for durable goods.

III. International Evidence on Consumption Variability

A. Estimation of φ_n and φ_d

Given a real return on nonhuman wealth close enough to zero, the results in the lemma above suggest the use of the following estimators:

$$\hat{\varphi}_n \equiv 1/\sqrt{2\pi \hat{f}_n(0)}$$

$$\hat{\varphi}_d \equiv \mu(\delta)/\sqrt{2\pi \hat{f}_d(0)}$$

where $\hat{f}_n(0)$ and $\hat{f}_d(0)$ are consistent estimators for the normalized spectrum of Δcn and Δcd at frequency zero.

The results reported below are based on the Bartlett estimator,[8] defined as

$$\hat{f}_i(0) = (1/2\pi)\left\{1 + 2\sum_{s=1}^{M}[1-(s/M)]\hat{\rho}_i(s)\right\} \qquad i = n, d$$

where $\hat{\rho}_i(s)$ is a consistent estimator of the autocorrelogram of either Δcn or Δcd at lag s, and M is the parameter controlling the window size. Consistency of the above estimator requires that $M \to \infty$ and $(M/N) \to 0$, as $N \to \infty$, where N is the number of observations available. In practice, however, N is given and a finite value of M has to be chosen, which involves a trade-off between the bias and variance of $\hat{f}_i(0)$. Below we report estimates based on different M values, in order to check the robustness of the results.

B. Data

Our data come from two different sources. Data for the United States were obtained from Citibase, and cover the period 1947:1–1989:4. Data for five other countries (Canada (1960:1–1989:4), United Kingdom (1963:1–1989:4), Japan (1970:1–1989:4), France (1970:1–1989:4) and Italy (1970:1–1989:4)), were taken from OECD's *Quarterly National Accounts* publication, table 6b. Our choice of countries and sample periods was made on the basis of data availability. For each country two consumption measures are used, *cd* corresponds to consumption expenditures on durable goods. *cn* is the difference between total private consumption expenditures and *cd*, and thus includes both nondurables and services. Per capita series were constructed using quarterly data on population size. The corresponding data for the United States were obtained from Citibase. Quarterly measures of population size for the remaining countries were constructed by interpolating between annual mid-year estimates taken from IMF's *International Financial Statistics*. With the exception of Japan, all data were already seasonally adjusted. We seasonally adjusted Japanese data by regressing Δcn and Δcd on seasonal dummies, allowing for a time trend in the seasonal when significant.

All the time series finally used are expressed in constant domestic prices, annual rates, seasonally adjusted, and on a per capita basis.

C. Preliminary Analysis: Unit Roots and Cointegration Tests

Implicit in the analysis above is the assumption that both *cn* and *cd* are $I(1)$ processes, i.e., processes stochastically integrated of order one.[9] The unit root in consumption has some theoreti-

[8] For a detailed discussion of the Bartlett estimator, see Priestley (1981)

[9] That assumption is implicit in (13) and (14), combined with our assumption of continuity of $\beta_i(\lambda)$ and $\phi_i(\lambda)$ at $\lambda = 1$.

TABLE 1.—UNIT ROOT TESTS

	United States	Canada	United Kingdom	Japan	Italy	France
Nondurable, levels						
ADF[a]	−2.73	−1.78	0.88	−1.78	-3.86^{c}	−2.57
Z[b]	−2 38	−1 94	0 74	−2.44	−2.15	−2 74
Durables, levels						
ADF	−1.36	−1.76	−1.70	−2.55	−1.39	−2.15
Z	−1.57	−2.01	−2.28	−2.09	−0.82	−2 37
Nondurables, differences						
ADF	-7.70^{c}	-6.83^{c}	-5.79^{c}	-7.39^{c}	-3.33^{d}	-7.88^{c}
Z	$-11\ 06^{c}$	-10.37^{c}	-10.16^{c}	-10.96^{c}	-3.33^{d}	-12.54^{c}
Durables, differences						
ADF	-9.02^{c}	$-6\ 48^{c}$	-14.83^{c}	$-7\ 36^{c}$	$-3\ 66^{c}$	-6.08^{c}
Z	-14.31^{c}	$-11\ 40^{c}$	-15.65^{c}	$-9\ 12^{c}$	$-4\ 84^{c}$	$-8\ 62^{c}$

[a] Augmented Dickey-Fuller (1979) t-statistic, based on an OLS regression of each variable on its own lag, its lagged first-difference, a time trend, and an intercept Asymptotic critical values are −3 12 (10%) and −3 41 (5%)

[b] Phillips-Perron (1988) t-statistic, based on an OLS regression of each variable on its own lag, a time trend, and an intercept Asymptotic critical values are −3 12 (10%) and −3 41 (5%)

[c] Significant at the 5% level

[d] Significant at the 10% level

cal appeal: it arises in a variety of models as a result of households' willingness to smooth consumption over time, *independently* of the order of integration of the labor income process.[10]

Table 1 reports the results of unit root tests based on ADF (Dickey and Fuller (1979)) and Phillips-Perron (1988) t-statistics applied to the data described above. The tests systematically fail to reject the unit root null against the trend-stationary alternative at conventional significance levels. The only exception to that result is given by the ADF test applied to nondurables consumption in Italy, which rejects the null of a unit root for that series at the 5% significance level. When similar tests are applied to first-differences Δcn and Δcd, all the statistics reject the unit root null at low significance levels.

The long-run identifying restriction (19) implies that, in the absence of preference shocks, cn and cd should be cointegrated. In the context of our model, lack of cointegration between the two variables may still result from the presence of a unit root in the "preference" components of cn and cd. However, as shown above, and given sufficiently low interest rates, the requirement that the intertemporal budget constraint is met implies that $\phi_i(1) \cong 0$, for $i = n, d$. Accordingly, if both our identifying assumption and our approximation are correct we should expect cn and cd to be nearly cointegrated.

Table 2 reports the results of several cointegration tests. The first set of statistics is based on the residuals of an OLS regression of cd on cn, a time trend and an intercept (regression #1), the second set being based on the residuals of the "inverted" regression (regression #2). We report both the ADF and Phillips t-statistics associated with the null of a unit root in the residual of the cointegrating regression (see Phillips and Ouliaris (1990)). For each country, we apply each statistic to data covering the full sample period, as well as a truncated sample excluding data of the eighties.

The results based on the full sample are far from overwhelming. Using the residuals from the first regression, cointegration can only be established at conventional significance levels for the United Kingdom and France, and, in the latter case, only with the Phillips t-statistic, at the 10% level. The tests based on the second regression suggest that cointegration may be present also in Japan and the United States, but in both cases that result is supported by only one of the two statistics only. When we exclude the eighties from our sample period the results become much stronger. Cointegration can be established with at least one statistic for all countries considered, at a 5% significance. Notice that the test results are not substantially affected by the cointegrating regression used with the exception of Italy. Given the previous results, we choose to present sepa-

[10] There are exceptions, however, including RBC models (e.g , King, Plosser and Rebelo (1988)) and overlapping generations models (e.g. Galí (1990)), when the driving technology process is stationary.

TABLE 2.—COINTEGRATION TESTS

	United States	Canada	United Kingdom	Japan	Italy	France
CI Regression #1[a]						
Full sample						
ADF[b]	−2.94	−1.71	−3.81[e]	−2.96	−0.86	−2.67
Z[c]	−2.98	−1.97	−5.79[e]	−2.95	−0.55	−3.57[f]
Truncated sample						
ADF	−4.11[e]	−3.03	−3.71[f]	−4.67[e]	−2.14	−2.97
Z	−4.12[e]	−3.93[e]	−5.81[e]	−4.46[e]	−1.58	−4.24[e]
CI Regression #2[d]						
Full sample						
ADF	−3.73[f]	−1.70	−2.60	−2.43	−2.83	2.67
Z	−3.46	−2.22	−5.51[e]	−5.22[e]	−2.79	−6.21[e]
Truncated sample						
ADF	−4.30[e]	−3.10	−2.51	−3.00	−3.87[e]	−3.84[e]
Z	−4.54[e]	−5.18[e]	−5.81[e]	−9.07[e]	−4.61[e]	−12.81[e]

[a] Tests based on the residual of an OLS regression of cd on cn, a time trend, and an intercept

[b] Augmented Dickey-Fuller t-statistic corresponding to a unit root test on the residual of the cointegrating regression Asymptotic critical values are −3.51 (10%) and −3.80 (5%) See Phillips and Ouliaris (1990)

[c] Phillips t-statistic corresponding to a unit root test on the residual of the cointegrating regression Asymptotic critical values are −3.51 (10%) and −3.80 (5%) See Phillips and Ouliaris (1990)

[d] Tests based on the residual of an OLS regression of cn on cd, a time trend, and an intercept

[e] Significant at the 5% level

[f] Significant at the 10% level

TABLE 3.—NONDURABLES VARIABILITY RATIOS

M	United States	Canada	United Kingdom	Japan	Italy	France
Full sample						
5	0.749	0.834	0.775	1.145	0.563	1.274
	(< 0.01)	(< 0.01)	(< 0.01)	(0.14)	(< 0.01)	(0.05)
10	0.679	0.726	0.642	1.158	0.560	1.294
	(< 0.01)	(< 0.01)	(< 0.01)	(0.27)	(< 0.01)	(0.10)
20	0.637	0.636	0.586	1.161	0.656	1.187
	(< 0.01)	(< 0.01)	(< 0.01)	(0.30)	(< 0.01)	(0.22)
Truncated sample						
5	0.767	0.928	0.954	1.175	0.573	1.426
	(< 0.01)	(0.23)	(0.31)	(0.22)	(< 0.01)	(0.05)
10	0.725	0.821	0.885	1.210	0.654	1.504
	(< 0.01)	(0.06)	(0.18)	(0.24)	(< 0.01)	(0.03)
20	0.642	0.704	1.118	1.207	0.675	1.373
	(< 0.01)	(0.03)	(0.68)	(0.27)	(0.02)	(0.04)

Note Figures reported are the estimates of variability ratio φ_n, using a Bartlett-window of size M Figures in brackets are p values for the H_0 $\varphi_n = 1$ null See description of estimator and data in main text

rate evidence for both the full sample and the truncated sample excluding the eighties in the remainder of the paper.

D. *Empirical Results: Variability Ratios for Nondurables*

Table 3 reports the estimates of variability ratio φ_n for each of the six countries considered. For each country and sample period we report the estimates corresponding to three different values of the window size M. In brackets, we report the p-values corresponding to a one-sided test of the null hypothesis H_0: $\varphi_n = 1$, against the relevant alternative.[11] The use of asymptotic distribution in such tests is potentially misleading since, among other reasons, $\hat{f}_n(0)$-type estimators, though consistent, are biased in small samples.[12] Consequently, we report simulated p-values, ob-

[11] That is, H_A $\varphi_n < 1$ if $\hat{\varphi}_n$ is less than one, H_A. $\varphi_n > 1$ if $\hat{\varphi}_n$ is greater than one.

[12] See Galí (1991) for more on this issue

TABLE 4.—DURABLES VARIABILITY RATIOS

M	United States	Canada	United Kingdom	Japan	Italy	France
Full sample						
5	0 037	0.034	0 052	0 041	0 020	0.033
	(< 0.01)	(< 0.01)	(< 0.01)	(< 0 01)	(< 0 01)	(< 0.01)
10	0.036	0.033	0.056	0.041	0.018	0 035
	(< 0.01)	(< 0.01)	(< 0.01)	(< 0.01)	(< 0.01)	(< 0 01)
20	0.037	0.033	0.062	0 042	0 017	0.040
	(< 0.01)	(< 0 01)	(< 0.01)	(< 0 01)	(< 0 01)	(< 0.01)
Truncated sample						
5	0.028	0.034	0 035	0.045	0.021	0.057
	(< 0.01)	(< 0.01)	(< 0 01)	(< 0.01)	(< 0.01)	(< 0.01)
10	0.026	0 030	0 033	0 052	0.024	0 075
	(< 0.01)	(< 0.01)	(< 0 01)	(< 0.01)	(< 0.01)	(< 0 01)
20	0.025	0.028	0 037	0.058	0.026	0 083
	(< 0 01)	(< 0 01)	(< 0.01)	(< 0.01)	(< 0.01)	(< 0 01)

Note Figures reported are the estimates of variability ratio φ_d, using a Bartlett-window of size M Figures in brackets are p values for the H_0 $\varphi_d = 1$ null See description of estimator and data in main text

tained from the empirical distribution of the estimator $\hat{\varphi}_n$ under the PIH null, given the relevant sample size.[13]

The point estimates of $\hat{\varphi}_n$ are significantly less than one in the United States, Canada, United Kingdom and Italy, and their values are not much affected by the window size. Since those estimates are consistent for the "true" φ_n ratio (despite the rejection of the PIH null), the results imply that consumption is *too smooth* in those countries, relative to the prediction of the permanent income model. That excess smoothness finding is strongly significant for the four countries when data for the full sample are used, but is somewhat weakened for Canada and the United Kingdom when we use the truncated sample.

For both Japan and France, the $\hat{\varphi}_n$ estimates are greater than one for all window sizes, but are only significantly so (for most window sizes) for France. The results thus point toward evidence of excess volatility in nondurables consumption in the latter country. The variability of nondurables consumption in Japan appears to be in the range predicted by the permanent income model.

E. Empirical Results: Variability Ratios for Durables

Table 4 reports the estimates of variability ratio φ_d, denoted by $\hat{\varphi}_d$. The estimates reported are conditional on a value of 0.05 for δ, which roughly corresponds to a 20% annual depreciation rate. The previous number is the average rate implicit in the gross expenditure on durables and durables net stock series in the United States, as reported in Bernanke (1985). The p value below each estimate was obtained under the PIH null, which corresponds in this case to the MA(1) process for Δcd in (8). The results are quite strong. All countries show extremely low $\hat{\varphi}_d$ estimates, with values ranging from 0.02 (Italy) to 0.06 (Japan). In all cases considered the $\varphi_d = 1$ null can be rejected at very low significance levels. The results thus suggest the presence of *extreme excess smoothness* in the time series for durables consumption, whose first difference is for all countries considered about 20 times less variable than predicted by the permanent income model.

Even though our estimator $\hat{\varphi}_d$ is consistent for the true variability ratio (conditional on δ) there are good reasons to believe it may be subject to a serious bias problem in small samples, both under the null and under alternatives "close" to the null. In addition, the use of a "wrong" δ value may generate quantitatively misleading estimates, even asymptotically.[14] Interestingly, it is possible to assess the relative variability of durables consumption in a way which does not require obtaining a (possibly highly biased) estimate of $f_d(0)$. To see this, notice that *under the permanent in-*

[13] In a previous version of this paper tests of a continuous time version of the PIH were also performed and reported None of the conclusions were substantially altered by that modification.

[14] Let δ^* and δ denote the true and assumed depreciation rates, respectively. For plausibly low values of both parameters, $\lim_{N \to \infty} \hat{\varphi}_d \cong (\delta/\delta^*)\varphi_d$.

TABLE 5.—DURABLES VS. NONDURABLES VARIABILITY

	United States	Canada	United Kingdom	Japan	Italy	France
Full sample						
$\hat{\Omega}_*$ estimates[a]						
$\delta = 0.05$	18.75	11.49	8.83	8.78	17 37	14.79
$\delta = 0.025$	38.44	23.56	18.11	18.01	35 62	30.34
$\delta = 0.01$	8.91	5.46	4.20	4.17	8.26	7.03
$\hat{\Omega}$ estimates[b]	0.25	0.26	0.02	0 15	0.26	0.17
Truncated sample						
$\hat{\Omega}_*$ estimates						
$\delta = 0.05$	13.47	10 72	12.17	5.62	19.00	12 16
$\delta = 0.025$	27.63	21 99	24.95	11.52	38 97	24 95
$\delta = 0.10$	6.41	5.10	5.78	2.67	9.04	5.78
$\hat{\Omega}$ estimates	0.21	0.24	0.18	0.20	0.21	0.21

[a] Estimate of $\Omega_* \equiv \sigma_*(\Delta cd/\sigma_*(\Delta cn)$, i e , the ratio of durables/nondurables standard deviations implied by the PIH model
[b] Estimate of $\Omega \equiv \sigma(\Delta cd/\sigma(\Delta cn)$, i e , the actual ratio of durables/nondurables standard deviations

come model developed above the standard deviation of Δcd relative to that of Δcn (denoted by Ω_*) is given by (see (1)–(12)):

$$\Omega_* \equiv \sigma_*(\Delta cd)/\sigma_*(\Delta cn) = [\delta\gamma/\theta]/\mu(\delta). \tag{24}$$

Under the maintained assumption of cointegration between *cn* and *cd* (see (9)), consistent estimates of $[\delta\gamma/\theta]$ for each country were obtained in the cointegrating regression of *cd* on *cn*. The first three rows in table 5 show the estimate of Ω_* (denoted by $\hat{\Omega}_*$) implied by (24), given the cointegrating vector estimates and $\mu(\delta)$ values consistent with alternative assumptions on δ. Let Ω denote the *actual* variability ratio $\sigma(\Delta cd)/\sigma(\Delta cn)$. The fourth row in table 5 reports a consistent estimate of Ω, denoted by $\hat{\Omega}$, equal to the ratio of sample standard deviations. Clearly, the difference between the estimates of predicted ($\hat{\Omega}_*$) and actual ($\hat{\Omega}$) ratios is very large for all countries and depreciation rates considered.

Since we can always express $\varphi_d \equiv \sigma(\Delta c^d)/\sigma_*(\Delta c^d)$ as

$$\varphi_d \equiv \varphi_n \Omega/\Omega_*$$

and since the evidence in the previous subsection pointed to values of φ_n *relatively* close to one, the low values of Ω/Ω_* implicit in table 5 are an additional manifestation of the extreme excess smoothness shown by durables consumption relative to permanent income innovations. Similar results are obtained when the truncated sample is used, as reported in the second half of table 5.

F. *Empirical Results: Discussion*

The results in the previous subsection can be analyzed from two alternative angles: we can focus on the observed patterns *across goods*, or, alternatively, on the observed patterns *across countries*.

Patterns across Goods: Our results provide strong evidence against the permanent income model for both durables and nondurables consumption. Nevertheless, an important asymmetry between the two types of consumption seems to emerge. The rejection of the PIH in the durables case is much stronger, with *all* the estimates pointing towards the presence of excess smoothness for all countries. On the other hand, rejections of the PIH for nondurables, though applying more or less strongly to all the countries considered, are quantitatively small, and have different signs for different countries.

That asymmetry in the size of variability ratios across goods, if confirmed by further work, should prove useful in the process of identifying the sources of rejection of the permanent income model. From that perspective, explanations of excess smoothness that rely on habit formation (e.g., Deaton (1987)) would require more "habit persistence" on the services of durable goods than on nondurables consumption, a requirement on preferences that does not seem, at first thought, particularly intuitive. Thus, the observed differences in variability ratios across goods tend to rule out an explanation based on habit formation.

On the other hand, the existence of significant transaction costs (information gathering, illiquid secondary markets, etc.) in the adjustment of the stock of durables (likely to be little relevant for nondurables purchases) may underlie a more promising explanation. Excess smoothness in durables purchases arises naturally in models with either quadratic adjustment costs and "smooth," continuous adjustment at the micro level (e.g., Bernanke (1985)) or in more complex models with fixed costs and "lumpy" adjustment at the micro level, but sluggishness in the aggregate (e.g., Bertola and Caballero (1990)).

Liquidity constraints and/or myopia may also be consistent with the observed differences between the relative variability of durables and nondurables consumption. If liquidity constrained consumers follow a rule of thumb that consists in spending a constant fraction of their income on durables and the rest on nondurables, the time series properties for their Δcn and Δcd will resemble those of income changes. If income follows a random walk, Δcn will show no serial correlation, regardless of the relative importance of liquidity constrained consumers, and thus behave as if the PIH held. On the other hand, and to the extent that liquidity constrained consumers account for a significant fraction of aggregate consumption, Δcd will fail to show the large, negative MA coefficient found in (8), and thus exhibit excess smoothness.

Patterns across Countries: If the liquidity constraints hypothesis is correct, departures from the PIH model should be larger in economies with poorly developed credit markets. Jappelli and Pagano (1989) and Campbell and Mankiw (1991) looked at the cross-country evidence on *predictability* of consumption, and concluded that the evidence was roughly consistent with that hypothesis. Here we look instead at the estimated variability ratios.

On the basis of the size of the market for consumer debt (Jappelli and Pagano (1989)) the United States is likely to have the most developed credit markets, followed by the United Kingdom and Japan, both of which are in turn ahead of Italy. Jappelli and Pagano do not provide any information for Canada and France, but Campbell and Mankiw (1991) suggest Canada and France should be classified as having well developed and relatively undeveloped credit markets, respectively.

Perhaps surprisingly, no obvious correlation emerges between estimated variability ratios for different countries and their "relative position" in the informal ranking above. The estimated variability ratios for durables consumption, take extremely low and not too different values for all countries, without any clear correlation between those values and the credit market ranking. For nondurables the lack of correlation takes a different form: strong excess smoothness is detected in countries with both well-developed credit markets (United States), and undeveloped ones (Italy). On the other hand, the strongest evidence of excess volatility can be found in France, where consumers' access to credit markets is probably not too different from that in Italy.

Overall, and despite the informality of the previous discussion, it seems safe to conclude that the international pattern of results above does not provide further evidence supporting the liquidity constraints hypothesis as the main source of PIH rejections.

IV. Summary and Conclusions

The present paper has extended the econometric approach to identification and estimation of variability ratios developed in Galí (1991) to the case in which both durables and nondurables consumption coexist. That approach avoids some of the shortcomings of other methods found in the literature. In particular, our method does not impose any restriction on the stochastic process for labor income nor on the consumers' information set. The estimators developed are consistent under relatively unrestricted short-run departures from the permanent income model.

Using our method, we estimate consumption variability ratios for both durables and nondurables consumption, using data for six OECD countries. Some important departures from the permanent income model emerge: (i) nondurables consumption measures show mild excess smoothness in the United States, Canada, United Kingdom, and Italy, and mild excess volatility in

France, and (ii) durables consumption measures show extreme excess smoothness in all countries.

Those results seem consistent with the presence of significant adjustment costs and/or liquidity constraints. Given their different implications for policy, further research aimed at sorting out their relative importance as sources of excess smoothness is needed. That work will require looking at alternative models' predictions other than those which can be expressed in terms of the spectrum of consumption changes, and will thus have to involve tools and information beyond those used in the present paper.

APPENDIX

Proof of Lemma

Given (12), the spectrum of Δcn at frequency zero is given by

$$h_n(0) = (1/2\pi)\{\beta_n(1)^2\sigma^2(\xi) + \phi_n(1)^2\sigma^2(\eta)\}.$$

Let $f_n(\omega) \equiv h_n(\omega)/\sigma^2(\Delta cn)$ denote the normalized spectrum of Δcn. It follows that

$$\{1/\sqrt{2\pi f_n(0)}\} = \sigma(\Delta cn)/\sqrt{\beta_n(1)^2\sigma^2(\xi) + \phi_n(1)^2\sigma^2(\eta)}$$

Taking the limit on a sequence of economies for which $r \to 0$, and using (15) and (16), we get

$$\lim_{r\to 0}\left[\{1/\sqrt{2\pi f_n(0)}\} - \{\sigma(\Delta cn)/\beta_n(1)\sigma(\xi)\}\right] = 0$$

and (22) in the lemma follows, for (11) and (20) imply

$$\lim_{r\to 0}\left[\{\sigma(\Delta cn)/\beta_n(1)\sigma(\xi)\} - \varphi_n\right] = 0$$

The durables case is handled symmetrically The spectrum of Δcd is given by

$$h_d(0) = (1/2\pi)\{\beta_d(1)^2\sigma^2(\xi) + \phi_d(1)^2\sigma^2(\nu)\}.$$

It follows that

$$\mu(\delta)/\sqrt{2\pi f_n(0)} = \mu(\delta)\sigma(\Delta cd)/\sqrt{\beta_d(1)^2\sigma^2(\xi) + \phi_d(1)^2\sigma^2(\nu)}.$$

Taking the limit as $r \to 0$, and using (15) and (16), we get

$$\lim_{r\to 0}\left[\{\mu(\delta)/\sqrt{2\pi f_d(0)}\} - \{\mu(\delta)\sigma(\Delta cd)/\beta_d(1)\sigma(\xi)\}\right] = 0.$$

From (12) and (21) we know

$$\lim_{r\to 0}\left[\{\mu(\delta)\sigma(\Delta cd)/\beta_d(1)\sigma(\xi)\} - \varphi_d\right],$$

which, combined with the result above, yields (23).

REFERENCES

Bernanke, Ben, "Adjustment Costs, Durables, and Aggregate Consumption," *Journal of Monetary Economics* 15 (1985), 41–68

Bertola, Giuseppe, and Ricardo J Caballero, "Kinked Adjustment Costs and Aggregate Dynamics," in Olivier J. Blanchard and Stanley Fischer (eds), *NBER Macroeconomics Annual 1990* (Cambridge, MA· M I.T Press, 1990)

Campbell, John Y , and Angus Deaton, "Why Is Consumption so Smooth?" *Review of Economic Studies* 56 (July 1989), 357–374.

Campbell, John Y , and N. Gregory Mankiw, "Consumption, Income and Interest Rates Reinterpreting the Time Series Evidence," in Olivier J. Blanchard and Stanley Fischer (eds), *NBER Macroeconomics Annual 1989* (Cambridge, MA: M.I.T Press, 1989)

______, "The Response of Consumption to Income· A Cross-Country Investigation," *European Economic Review* (1991).

Clarida, Richard H , "Smooth Consumption and Noisy Permanent Income Is There a Puzzle?," Columbia University mimeo, June 1988.

Deaton, Angus, "Life-Cycle Models of Consumption Is the Evidence Consistent with the Theory?" in T Bewley (ed), *Advances in Econometrics, Fifth World Congress*, Vol. 2 (Cambridge. Cambridge University Press, 1987).

Dickey, D A , and W. A. Fuller, "Distribution of the Estimators for Autoregressive Time Series with a Unit Root," *Journal of the American Statistical Association* 74 (1979), 427–431

Flavin, Marjorie, "The Adjustment of Consumption to Changing Expectations about Future Income," *Journal of Political Economy* 89 (Oct 1981), 974–1009

______, "The Excess Smoothness of Consumption. Identification and Interpretation," NBER Working Paper #2807 (Dec 1988)

Galí, Jordi, "Finite Horizons, Life Cycle Savings and Time Series Evidence on Consumption," *Journal of Monetary Economics* 26 (Dec. 1990), 433–452

______, "Budget Constraints and Time Series Evidence on Consumption," *American Economic Review* 81 (Dec 1991), 1238–1253.

Hall, Robert E , "Stochastic Implications of the Life-Cycle-Permanent-Income Hypothesis Theory and Evidence," *Journal of Political Economy* 86 (Dec 1978), 971–987

Jappelli, Tullio, and Marco Pagano, "Consumption and Capital Market Imperfections An International Comparison," *American Economic Review* 79 (Dec 1989), 1088–1105

Mankiw, N. Gregory, "Hall's Consumption Hypothesis and Durable Goods," *Journal of Monetary Economics* 10 (1982), 417–425

Phillips, P. C. B , and S. Ouliaris, "Asymptotic Properties of Residual Based Tests for Cointegration," *Econometrica* 58 (1990), 165–193.

Phillips, P C. B , and Pierre Perron, "Testing for a Unit Root in Time Series Regression," *Biometrika* 75 (1988), 335–346.

Priestley, M. B , *Spectral Analysis and Time Series* (San Diego Academic Press, 1981)

Quah, Danny, "Permanent and Transitory Movements in Labor Income· An Explanation for Excess Smoothness in Consumption," *Journal of Political Economy* 98 (June 1990), 449–475.

West, Kenneth D , "The Insensitivity of Consumption to News about Income," *Journal of Monetary Economics* 21 (Jan 1988), 17–34.