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THE PERMANENT INCOME HYPOTHESIS WHEN THE BLISS POINT IS STOCHASTIC

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ABSTRACT

A version of the permanent income model is developed in which the bliss point of the agent is stochastic. The bliss point depends on realizations of the stochastic process generating labor income and a random shock. The model predicts consumption and labor income share a common trend and that a linear combination of current consumption, current labor income, and once lagged consumption is stationary. Empirically, consumption appears more serially correlated than the model is capable of supporting. Further, the volatility of consumption appears sensitive to time variation in real interest rates.

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I. Introduction

A stylized fact of macroeconomics is the ability of economists to create paradoxes for themselves within the context of long accepted theories. The Deaton paradox and what can be called West's corollary fit this description. The paradox of Deaton (1988), as interpreted by Deaton and Campbell (1989), is that although permanent labor income is a 'noisy' time series, aggregate consumption is observed to be smoother than disposable income. As a result, the standard permanent income hypothesis (PIH) predicts the time series of consumption should be more variable than the time series of labor income. The empirical fact for aggregate post-war U.S. data is the opposite.

The corollary of West (1988b), given labor income possesses a unit root, is that empirically consumption is insensitivity to innovations in lagged labor income. This is even after making account for the inability of the econometrician to observe all the information agents in the economy use in forecasting future labor income. The upshot of West (1988b) is that given a unit root in labor income it appears an empirical regularity is that news about labor income has no implications for consumption.

To reconcile the excess smoothness quandary, Quah (1990) distinguishes between permanent and transitory innovations in labor income. Once the univariate process for labor income is broken into these two components, the properties of the univariate process generating labor income are not particularly relevant for the predictions of the standard PIH concerning the smoothness of consumption. However, as Quah (1990) notes this leaves the observed non-martingale behavior of aggregate consumption unexplained.

The martingale prediction of the standard PIH, that the best forecast of future consumption is current consumption, is the other key restriction on the data made by the standard PIH. Empirically this prediction of the standard PIH has been found wanting too. The results of Hall (1978), Flavin (1981), Christiano (1987a) have in one way or another rejected the martingale restriction.[1] These results have come to be known as the excess sensitivity debate while the results of Deaton and Campbell (1989) and West (1988b) have come to be known as the excess smoothness quandary. Although the excess smoothness quandary relies on a unit root in labor income, the excess sensitivity debate does not.

The purpose of this paper is to address some of the empirical issues concerning the PIH within an economy in which the preferences or tastes of agents change through time and labor income possesses a unit root. The model described below follows the permanent income model found in Hansen (1988) and Hansen and Sargent (1990). This version of the permanent income model is a dynamic economy consisting of an infinitely lived representative agent who possess quadratic utility, has access to a linear production technology, and a complete set of state-contingent claims markets. One unique feature of this economy is the presence of a preference or taste variable in the utility function of the representative agent which is a linear function of the forcing variables of the economy.[2] The manner in which the taste variable enters the utility function of the agent means it acts as a stochastic bliss point. An outcome of this is the long run optimal consumption plan of the representative agent is to set her consumption equal to her stochastic bliss point.

The parametric form of the stochastic bliss point adopted here along with

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[1] Nelson (1987) points out that the results in Flavin (1981) may be the result of inappropriate detrending procedures.

[2] Equilibrium models of time-varying or changing tastes have been studied for over twenty years. Some relevant models of micro behavior are Pollak (1976a, 1976b) and Boyer (1983). Models of aggregate behavior include Ryder and Heal (1973), Boyer (1978), Sundaresan (1989), Eichenbaum and Hansen (1990), Constantinides (1990), Abel (1990), Backus, Gregory, and Telmer (1990), Heaton (1990, 1991), Ogaki and Park (1990), and Nason (1988, 1989).

the optimal consumption plan of the representative agent creates a set of predictions that differ from the standard PIH. For example, with a stochastic bliss point the PIH no longer predicts consumption is a martingale. Further, the PIH with a stochastic bliss point provides a set of identifying restrictions on the permanent and transitory shocks of the bivariate stochastic process of consumption and labor income. As in Quah (1990), these identifying restrictions provide a means for reconciling some of the empirical anomalies of consumption with a (version) of the PIH.

The plan of the paper follows. Sections II borrows liberally from Hansen (1988) and Hansen and Sargent (1990) to construct the permanent income model when the utility function possesses time-varying tastes. Section III discusses some econometric issues involved in cointegration and tests of the model of section II. In the context of the model of section II, section IV presents empirical results. A short conclusion appears as section V.

IIa. The Permanent Income Model with Changing Tastes

Hansen (1988) constructs a permanent income model with taste shocks building up from the optimization problems of finitely many heterogeneous consumers to the equilibrium laws-of-motion of the aggregate economy. Each agent has a period utility function of the form

$$U(c_t - \underline{c}_t) = -0.5(c_t - \underline{c}_t)^2, \quad c_t \le \underline{c}_t,$$

where c_t and $\underline{c}_t$ represent consumption and the taste variable or stochastic bliss point, respectively.[3] Hence, agents measure utility as the difference between their consumption and their stochastic bliss point and according to this utility function agents chase their stochastic bliss point with the level

[3] The standard PIH makes the bliss point constant. The empirical results in Lewbel (1987) suggest the quarterly U.S. aggregate time series rejects this specification of the bliss point.

of their consumption in order to maximize utility. It is possible to aggregate over the heterogeneous preferences of consumers into a *stand in* or representative agent because it is assumed the preferences of agents are quadratic and the market for contingent claims is complete.[4]

In order to generate testable implications for the PIH with a stochastic bliss point, the taste shock must be given explicit form. One simple way to do this, which is consistent with Hansen (1988), is to make the $\underline{c}_t$ common to all agents a linear function of the entire realized history of aggregate labor income, y_t, and a mean zero, independently and identically distributed (IID) random variable, u_t, with finite variance as in

$$\underline{c}_t = \underline{c}(y_t, \ldots, y_o) + u_t.$$

Idiosyncratic preference shocks can be thought of as being IID so that in the aggregate they wash out.[5]

The lifetime utility function of the infinitely lived representative agent (ILRA) is of the non-myopic exogenous changing taste form

$$E_0\left\{\sum_{t=0}^{\infty}\beta^t U(c_t - \underline{c}_t)\right\} \equiv -E_0\left\{\sum_{t=0}^{\infty}\beta^t\left[(c_t - \underline{c}_t)^2/2\right]\right\}, \quad 0 < \beta \equiv (1+\gamma)^{-1} < 1, \qquad (1)$$

where $0 < c_t < \underline{c}_t$, for all dates t.[6] The symbols $\mathbf{E}_0$ and β (γ) denote the expectations operator with respect to date zero information and the discount factor (the subjective rate-of-time preference), respectively. The utility function of the ILRA is non-myopiac and exogenous because the date t policy action of the ILRA has no effect on her utility during any date t+j, $j \geq 1$.

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[4] In this case, equilibrium quantities and prices can be interpreted as being generated in a competitive market economy in which the equilibrium distribution of wealth in the economy has no effect on the decision rules of the representative agent.

[5] Mace (1989) constructs and tests a model of full consumption insurance into which common and idiosyncratic preference shocks could be introduced.

[6] The term changing tastes is synonymous with the stochastic bliss point.

The date t information set of the ILRA is denoted by I_t, $E_t\{\cdot\} \equiv E\{\cdot/I_t\}$. It is assumed u_t is unknown to the econometrician. The information set available to the econometrician is denoted by H_t, $H_t \subset I_t$ ($u_{t-i} \notin H_t$, $i \geq 0$) and let $E^H_t\{\cdot\} \equiv E\{\cdot/H_t\}$.

To reflect the restrictions of the permanent income model, assume the ILRA owns a linear constant returns-to-scale storage technology with fixed gross return (1+r) for all dates t.[7] The budget constraint of the ILRA becomes

$$a_{t+1} = (1 + r)a_t + y_t - c_t, \tag{2}$$

where a_t is the physical wealth measured in units of the consumption good with which the agent enters date t, $r > 0$ is the constant rate-of-return on wealth between dates t and t+1, and r is measured per unit of the consumption good.

Existence of a stochastic competitive equilibrium for this economy has been shown by Hansen (1988) (among others). Key restrictions on a_t and γ in order for tractable solutions to be possible are noted with

$$\lim_{t\to\infty} (1 + \gamma)^{-t/2} a_t = 0.$$

This does not restrict $a_t \geq 0$, but means the ILRA cannot be made better off by holding physical capital indefinitely. The restriction on the limiting behavior of physical capital ensures a_{t+1} is measurable with respect to I_t.

The ILRA enters date t with a_t units of physical wealth which she invested in the linear production technology at the end of date t-1. Labor income and u_t are realized immediately after date t begins. Concurrently, the agent receives the return from her investment in the technology, $(1 + r)a_t$. Next, the ILRA makes her date t consumption (or saving) decision. This ends date t and has the ILRA entering date t+1 with a_{t+1} units of physical wealth.

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[7] Christiano (1987b) shows that a real business cycle model with a production technology linear in capital yields a rate-of-return equal across assets in the economy and constant for all time; also see Hansen and Sargent (1990).

IIb. The Decision Rules of the ILRA with Changing Tastes

In order to derive the predictions of the PIH with changing tastes, maximize (1) over uncertain consumption streams subject to (2), $0 \le c_t \le \underline{c}_t$, and given a_o. The first-order necessary conditions yield the Euler equation

$$-\left(c_t - \underline{c}_t\right) + \beta(1 + r)E_t\left\{c_{t+1} - \underline{c}_{t+1}\right\} = 0. \tag{3}$$

The PIH sets $\gamma = r$. Hansen (1988) shows that when this restriction is assumed

$$(1 - L)a_{t+1} = y_t - \underline{c}_t - \left(y_t^p - \underline{c}_t^p\right), \tag{4}$$

and

$$c_t = \underline{c}_t + ra_t + y_t^p - \underline{c}_t^p, \tag{5}$$

are the laws-of-motion for (aggregate) physical wealth and consumption, respectively, where $Lx_t = x_{t-1}$ and $x_t^p \equiv r(1+r)^{-1}\sum_{j=0}^{\infty}(1+r)^{-j}E_t x_{t+j}$. The first difference in consumption can be found by lagging (5) once, subtracting c_{t-1} from c_t, and substituting (4) to produce

$$(1 - L)c_t = (1 - L)\underline{c}_t + \Delta y_t^p - \Delta\underline{c}_t^p, \tag{6}$$

where $\Delta x_t^p \equiv r(1+r)^{-1}\sum_{j=0}^{\infty}(1+r)^{-j}\left[E_t x_{t+j} - E_{t-1}x_{t+j}\right]$.

Another set of restrictions for the PIH under changing tastes can be derived from (3). A solution which satisfies (3) is the consumption plan

$$c_t = \underline{c}_t \tag{7}$$

for all dates t.[8] In order to study (7) and its implications for the PIH with changing tastes, the stochastic bliss point is defined to be the sum of u_t and a geometric average of the entire history of realized labor income parameterized by the 'persistence' parameter θ and the 'trend' parameter ϕ

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[8]Hansen (1988) notes that $\gamma = r$ and the solution given by (7) results in an undesirable law of motion for aggregate physical wealth. In particular, $\gamma = r$ results in the solution given by (7) and the law of motion for physical wealth given by the intertemporal budget constraint (2); see Hansen (1988) especially his footnote 6. Nonetheless, the transversality condition is not violated by this solution; see Hansen and Sargent (1990).

$$\underline{c}_t = \phi(1 - \theta)\sum_{i=0}^{\infty}\theta^i y_{t-i} + u_t,$$

where $\phi \in [0,\infty)$ and $\theta \in [0,1)$. The specification $\underline{c}_t$ makes the stochastic bliss point resemble the PIH solution for consumption of Friedman (1957) as constructed by Muth (1960). More importantly, as will be discussed the analog to the 'transitory consumption shocks' of Flavin (1981) are serially correlated as well as correlated with permanent shocks to labor income.[9]

To show this and develop testable implications of the PIH with changing tastes, passing through the operator $(1-\theta L)$ causes (7) to become a first-order stochastic difference equation in consumption

$$c_t = \theta c_{t-1} + \phi(1 - \theta)y_t + u_t - \theta u_{t-1}. \tag{8}$$

Define $q_t \equiv c_t - \phi(1-\theta)y_t - \theta c_{t-1}$ and (8) can be written as

$$q_t = (1-L)c_t - \phi(1-\theta)(1-L)y_t - \theta u_{t-1} - E_{t-1}\left\{(1-L)c_t - \phi(1-\theta)(1-L)y_t\right\} \tag{9}$$

Given labor income possesses a unit root, equation (9) states that the PIH with changing tastes renders a linear combination of current consumption, current labor income, and lagged consumption stationary.

On the basis of I_{t-1}, q_t *is partially predictable for the ILRA. That is* $E_{t-1}q_t = \theta u_{t-1}$ and consumption is not a martingale. The partial predictability of q_t has implications for testing the PIH with changing tastes. To develop the testable implications of the PIH, begin with

$$\underline{c}_t^p \equiv \phi y_t^p + r\left[1+r-\theta\right]^{-1}\left(c_t - \phi(1-\theta)y_t\right) + r(1+r)^{-1}u_t$$

so that using (7), (5) becomes

$$c_t = \phi(1-\theta)y_t + (1+r-\theta)\left[a_t + (1-\phi)r^{-1}y_t^p - (1+r)^{-1}u_t\right] \tag{5'}$$

which resembles a Keynesian consumption function. Equation (9) and the first difference of equation (5') are used to produce

$$q_t = \left[(1+r-\theta)(1-\phi)r^{-1}\right]\Delta y_t^p - (1+r-\theta)(1+r)^{-1}u_t - \theta u_{t-1}. \tag{10}$$

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[9] Falk and Lee (1990) discuss the importance of reconciling the restrictions of the PIH of Friedman (1957) with the restrictions of the PIH of Hall (1978) and Flavin (1981).

Equation (10) expresses the short run equilibrium dynamics of the economy.

One way to understand these short run dynamics is to use equation (10) to generate the model of taste shocks due to Hayashi (1988). A version of the taste shock model of Hayashi (1988) is found by setting $\phi = 0$ and $\theta = 1$ in (10) to produce

$$(1-L)c_t = \Delta y^p_t - r(1+r)^{-1}u_t - u_{t-1}.$$

As noted by Hayashi (1988), the introduction of taste shocks causes consumption to lose the random walk property. Moreover, if u_{t-1} and y_{t-1} were correlated, a regression of $(1-L)c_t$ on $(1-L)y_{t-1}$ would yield a statistically significant slope coefficient.[10] Of course, (10) yields the same implications.

The Keynesian consumption function (5′) of the PIH with changing tastes can be made to resemble the consumption function studied by Campbell and Mankiw (1989, 1990). The relationship between disposable income, $yd_t \equiv y_t + ra_t$, and consumption means equation (5′) can be rewritten as

$$c_t = \phi(1 - \theta)yd_t + \varphi_t,$$

where

$$\varphi_t = \Big[1+r[1-\phi(1-\theta)]-\theta\Big]a_t + (1+r-\theta)\Big[(1-\phi)r^{-1}y^p_t - (1+r)^{-1}u_t\Big].$$

In this case, an instrumental variables regression of $(1-L)c_t$ on $(1-L)yd_t$ will yield a significant slope coefficient irrespective of the choice of instruments because $(1-L)a_t$ is a function of u_t. Hence, the PIH with changing tastes is observationally equivalent to the model of Campbell and Mankiw (1989, 1990), but the policy implications of the two models are disparate.

Although (10) yields $E_{t-1}q_t = -\theta u_{t-1}$, a problem for an econometrician

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[10] Hayashi points out the moving average structure in his permanent income model with taste shocks could be the result of measurement errors in consumption. The source of the measurement errors in consumption could be either temporal aggregation as in Heaton (1990) and Christiano, Eichenbaum, and Marshall (1991), durability as in Eichenbaum and Hansen (1990), or sampling errors as in Wilcox (1988).

attempting to test the time series implications of the PIH with changing tastes is that $u_{t-j} \notin H_t$, $j \geq 0$. Nonetheless, the econometrician can infer the unknown random disturbance u_t from the information set H_t. Define

$$\Delta y_t^H = r(1+r)^{-1}\sum_{j=0}^{\infty}(1+r)^{-j}\left[E_t^H y_{t+j} - E_{t-1}^H y_{t+j}\right]$$

and the econometrician equates common terms in (8) and (10) to produce

$$u_t^H = (1+r-\theta)(1-\phi)(1+r)r^{-1}\left[2(1+r)-\theta\right]^{-1}\Delta y_t^H.$$

Hence, for the econometrician equation (10) becomes

$$q_t^H = (1+r-\theta)(1-\phi)(1+r)r^{-1}\left[2(1+r)-\theta\right]^{-1}\left[\Delta y_t^H - \theta\Delta y_{t-1}^H\right] \qquad (10')$$

and (10') has the forecasting properties

$$E_{t-1}^H q_{t+j}^H = \begin{cases} -\Gamma\theta\Delta y_{t-1}^H, & j=0, \\ 0, & j\geq 1, \end{cases}$$

where $\Gamma \equiv (1+r-\theta)(1-\phi)(1+r)r^{-1}[2(1+r)-\theta]^{-1}$.[11]

Equation (10') can be used to develop a test of the sensitivity of consumption to news about labor income as in West (1988b). From proposition 1 of West (1988a), the variance inequality is derived as

$$E\left[q_t - E\{q_t/I_{t-1}\}\right]^2 \leq E\left[q_t^H - E_{t-1}\{q_t^H/H_{t-1}\}\right]^2.$$

Let y_t be generated by a stationary pth order autoregressive integrated, ARI(p,1), stochastic process

$$\rho(L)(1-L)y_t = e_t,$$

where $\rho(L) = 1 - \rho_1 L - \ldots - \rho_p L^p$, $E(e_t) = 0$, and $E(e_t^2) = \sigma_e^2$. Given this autoregression for $(1-L)y_t$, the formula of Hansen and Sargent (1980), and the results in West (1988a, 1988b), the variance inequality becomes

$$\sigma_q^2 \leq \Gamma^2(1+\theta^2)\rho([1+r]^{-1})^{-2}\sigma_e^2, \qquad (11)$$

where $\sigma_q^2 \equiv E\left[q_t - E\{q_t/I_{t-1}\}\right]^2$; see the appendix for details.

Equations (10') and (11) summarizes all the testable implications of the PIH with changing tastes. The forecasting properties of q_t^H given by (10') can

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[11] Equation (10') can be developed from a Kalman filter procedure as well.

be exploited in order to construct tests of various restrictions of the PIH with changing tastes. For instance, there should be no variables in H_{t-1} that help to explain q_{t+1}. However, note that tests of this prediction are weak restrictions of model. The variance bounds test of (11) contains the stronger restrictions of the PIH with changing tastes.

The solution given by (7), the optimality of setting consumption equal to the bliss point and altering a_{t+1} to satisfy this consumption plan provides a set of long run or low frequency restrictions on the bivariate stochastic process of consumption and labor income. The static or long run equilibrium of this economy as derived from (8) is

$$c_s = \phi y_s + u_s \tag{12}$$

where the subscript s denotes the static (long run) equilibrium value of a variable. The PIH with changing tastes (along with the parametric form of $\underline{c}_t$) makes $c_s - \phi y_s$ stationary. For physical wealth, the long run equilibrium relationship is found using the budget constraint (2) and equation (12)

$$a_s = -r^{-1}(1 - \phi)y_s - r^{-1}u_s$$

A nonstochastic long run equilibrium has $0 < a_s$ for $1 < \phi$.[12]

Equations (10′) and (12) have the virtues that the former predicts q_t is a stationary time series describing the short run dynamics of the model while the latter predicts the long run equilibrium relationship between consumption and labor income is stationary. These are the identifying restrictions on the permanent and transitory shocks of the PIH with changing tastes which are observed by the econometrician. However, the PIH with changing tastes presented here is only one of many possible identification schemes.[13]

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[12] The long run equilibrium solution of physical wealth can be found also by substituting equation (5′) into the intertemporal budget constraint (2) and solving for the long run equilibrium value of physical wealth.

[13] Quah (1991) identifies permanent and transitory shocks in labor income which

The long run restriction of the model is the bivariate relation between consumption and labor income given by (12). Given labor income possess a unit root, consumption and labor income share a common stochastic trend which has come to be know as a cointegrating relation. Let $\alpha' = [1\ -\phi]$, $z_t' = [c_t\ y_t]$ and $\eta_t = \alpha' z_t$ where α is the cointegrating vector and the trend parameter ϕ is the cointegrating parameter. The importance of cointegration for the estimation and testing of the PIH with changing tastes is pursued in the next section.

III. Cointegration, Structural Stability, and the PIH with Changing Tastes

The previous section presented the testable implications of the PIH with changing tastes. The econometric properties of the cointegrating relation, $\eta_t = \alpha' z_t$, depends on the order of integration of consumption and labor income. Given consumption and labor income are integrated of order one, I(1), the PIH with changing tastes makes a linear combination of consumption and labor income a stochastic process integrated of order zero, I(0); see Stock (1987).

Standard practice for estimating cointegrating relationships is to perform an ordinary least squares (OLS) regression of the type

$$c_t = \phi y_t + \eta_t. \tag{13}$$

With sample size T, Stock (1987) shows that the standard error of ϕ converges at rate T making the OLS estimate of ϕ, $\hat{\phi}$, super-consistent. It is worthwhile noting that the reason it is not possible to regress c_t on y_t and c_{t-1} to estimate ϕ and θ as suggested by (8) is that (i) ϕ possesses a non-standard distribution and (ii) it can be shown that θ is distributed asymptotically

drive the bivariate stochastic process of labor income and consumption when they are first differenced. If labor income possesses a unit root, the PIH with changing tastes makes it incorrect to test its predictions with $(1-L)y_t$ and $(1-L)c_t$; *see* Hansen, Roberds, and Sargent (1987). Nevertheless, the critique Quah (1991) outlines remains valid, but for the bivariate process of $(1-L)y_t$ and q_{t+1}.

normal; see Gregory, Pagan, and Smith (1991).

Several tests under the null of no cointegration have been developed by Engle and Granger (1987), Phillips (1987), and Phillips and Perron (1988). Granger and Engle (1987) propose to estimate an augmented Dickey-Fuller (ADF) regression to construct a simple t-ratio, t_δ, to test for a unit root in the OLS residuals, $\hat{\eta}_t$, from the regression of c_t on y_t. This is the OLS regression

$$(1-L)\hat{\eta}_t = \delta\hat{\eta}_{t-1} + \sum_{i=1}^{p}\nu_i(1-L)\hat{\eta}_{t-i} + w_t,$$

where the p lags of the first difference of the OLS residuals $\hat{\eta}_t$ are inserted to remove serial correlation in w_t.

The tests suggested by Phillips (1987) and Phillips and Perron (1988) require the regression

$$\hat{\eta}_t = \delta\hat{\eta}_{t-1} + w_t$$

to be estimated by OLS and calculation of the 'modified' t-ratio statistics

$$Z_\delta = T(\hat{\delta}-1) - 0.5(\hat{\omega}^2 - \hat{\sigma}_w^2)(T^{-2}\Sigma_{t=2}^{T}\hat{\eta}_{t-1}^2)^{-1}$$

and

$$Z_t = T(\hat{\delta}-1)/(\hat{\omega}^2\Sigma_{t=2}^{T}\hat{\eta}_{t-1}^2)^{1/2} - 0.5(\hat{\omega}^2 - \hat{\sigma}_w^2)(\hat{\omega}^2T^{-2}\Sigma_{t=2}^{T}\hat{\eta}_{t-1}^2)^{-1/2},$$

where $\hat{\sigma}_w^2 = T^{-1}\Sigma_{t=2}^{T}\hat{w}_{t-1}^2$ and $\hat{\omega}^2$ is an estimate of the variance of w_t at frequency zero. Estimation of $\hat{\omega}^2$ is discussed in Newey and West (1987). Critical values for t_δ, Z_δ, and Z_t are found in Phillips and Ouliaris (1990).[14]

A problem presented by the PIH with changing tastes is that η_t is serially correlated causing a bias in the tests for cointegration; see Gregory (1991). The source of the serial correlation of η_t is the persistence in consumption produced by θ. To show this, rewrite equation (10) as

$$(1-L)c_t = -(1-\theta)\left[c_{t-1} - \phi y_t\right] + v_t$$

where $v_t = [r^{-1}(1+r-\theta)(1-\phi)]\Delta y_t^p - (1+r-\theta)(1+r)^{-1}u_t - \theta u_{t-1}$ and substitute for

[14] The Monte Carlo evidence in Gregory (1991) suggests that the ADF t-ratio test of Granger and Engle (1987) and the Z_δ statistic of Phillips (1987) have reasonable small sample properties.

c_{t-1} in (10) to find

$$c_t = \phi y_t + (1-\theta)^{-1}\Big[v_t - \theta(1-L)c_t\Big], \quad \eta_t = (1-\theta)^{-1}\Big[v_t - \theta(1-L)c_t\Big].$$

Hence, η_t inherits the serial correlation present in consumption.

To work around the problem of the serial correlation of η_t. The first step is to estimate ϕ with OLS by regressing c_t on y_t. Although the tests for cointegration will be biased, the estimate of ϕ is super-consistent. The second step recognizes that the super-consistency of ϕ allows the uncertainty surrounding the OLS estimate of ϕ, $\hat{\phi}$, to be ignored when estimating the error correction equation

$$(1-L)c_t = -(1 - \theta)\alpha' z_{t-1} + (1 - \theta)(1-L)y_{\phi t} + v_t. \tag{14}$$

where $y_{\phi t} \equiv \hat{\phi} y_t$. Equation (14) follows directly from (10). To obtain a consistent estimate of θ, $\hat{\theta}$, (14) can be estimated by the generalized methods of moments (GMM) estimator of Hansen (1982).[15]

The process is repeated in a second round. Only now with a consistent estimate of θ the cointegration regression becomes

$$c_t^* = \phi y_t + v_t^*, \tag{15}$$

where $c_t^* = c_t + \hat{\theta}(1-\hat{\theta})^{-1}(1-L)c_t$ and $v_t^* = (1-\hat{\theta})^{-1}v_t$.[16] The transformation of the dependent variable into c_t^* removes the serial correlation in the error term and allows for unbiased tests of the null hypothesis of no cointegration. The final step of the second round is to use the second round estimate of ϕ in (14) to generate a second round GMM estimate of θ.

Besides OLS, two other methods are used to estimate ϕ as a check on the OLS estimate of ϕ. These are the fully-modified estimator of Phillips and Hansen (1990) and the GMM estimator of Hansen (1982) as discussed in West

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[15] The GMM estimator of Hansen (1982) is relevant because of the correlation between the first difference of labor income and v_t.

[16] The author wishes to thank Allan Gregory for suggesting the transformation.

(1988c). The Phillips and Hansen fully-modified (PH-FM) estimator and the GMM estimator of West (1988c) of cointegrating relations allow for inference on the cointegrating vector with classical asymptotic distribution theory.

PH-FM estimation involves applying OLS to (13) to construct $\hat{\eta}_t$. Next, $(1-L)y_t$ is demeaned and saved as $\hat{\zeta}_t$. Set $\hat{w}_t = [\hat{\eta}_t \;\hat{\zeta}_t]$ and adopt the Newey and West (1987) procedure to estimate the long run variance matrices

$$\hat{\Omega} = \sum_{j=-T}^{T} g_j T^{-1} \sum_{t=j+1}^{T} \hat{w}_{t-j}\hat{w}'_t,$$

and,

$$\hat{\Lambda} = \sum_{j=0}^{T} g_j T^{-1} \sum_{t=j+1}^{T} \hat{w}_{t-j}\hat{w}'_t,$$

where $g_j = 1 - |j/(T+j)|$. Partition $\hat{\Omega}$ and $\hat{\Lambda}$ by

$$\hat{\Omega} = \begin{bmatrix} \hat{\Omega}_{\eta\eta} & \hat{\Omega}_{\eta\zeta} \\ \hat{\Omega}_{\zeta\eta} & \hat{\Omega}_{\zeta\zeta} \end{bmatrix} \text{ and } \hat{\Lambda} = \begin{bmatrix} \hat{\Lambda}_{\eta\eta} & \hat{\Lambda}_{\eta\zeta} \\ \hat{\Lambda}_{\zeta\eta} & \hat{\Lambda}_{\zeta\zeta} \end{bmatrix}$$

to transform the dependent c_t with $c_t^+ = c_t - \hat{\Omega}_{\eta\zeta}\hat{\Omega}_{\zeta\zeta}^{-1}\hat{\zeta}_t$. The PH-FM estimator is

$$\hat{\phi}_{FM} = \left[\sum_{t=1}^{T}\left(c_t^+ y_t - T(0\;\hat{\Lambda}_{\eta\zeta}^{+\prime})\right)\right]\left[\sum_{t=1}^{T} y_t^2\right]^{-1},$$

where $\hat{\Lambda}_{\eta\zeta}^{+\prime} = \hat{\Lambda}_{\eta\zeta} - \hat{\Lambda}_{\zeta\zeta}\hat{\Omega}_{\zeta\zeta}^{-1}\hat{\Omega}_{\zeta\eta}$. The scores of the PH-FM estimator are defined by $\hat{\xi}_t = \left(y_t\hat{\eta}_t^+ - \begin{bmatrix} 0 \\ \hat{\Lambda}_{\zeta\eta}^+ \end{bmatrix}\right)$, where $\hat{\eta}_t^+ = c_t^+ - \hat{\phi}^+ y_t$ and $\sum_{t=1}^{T}\hat{\xi}_t = 0$.

The other check on the OLS estimate of ϕ is the estimator proposed by West (1988c). West (1988c) shows that if a set of I(1) variables which form a cointegrating relationship possess first differences with nonzero unconditional means, it is possible to estimate the cointegrating vector with GMM.[17] The GMM estimator of ϕ is

$$\hat{\phi}_{GMM} = \left[Y'G\tilde{S}^{-1}G'Y\right]^{-1} Y'G\tilde{S}^{-1}G'C,$$

where the tth rows of **Y** and **C** are y_t and c_t, **G** is a Txk matrix of (possibly nonstationary) instruments, and $\tilde{S}$ is the Newey and West (1987) corrected

[17] The proofs of asymptotic normality in West (1988c) rely on the regressors and error term having finite fourth moments too.

estimate of the variance of $\eta_t \otimes G_t$ at the zero frequency. To compute the variance of $\hat{\phi}_{GMM}$ construct

$$\hat{V}_{\phi GMM} = \left[Y'G\tilde{S}^{-1}G'Y\right]^{-1}$$

and $J(\phi) = \hat{e}G\tilde{S}^{-1}G'\hat{e}$ is the quadratic objective function of the GMM problem of (13). The $J(\phi)$ statistic is the Hansen (1982) test of instrument orthogonality and is distributed asymptotically χ^2 with k-1 degrees-of-freedom; see West (1988c) for details.

One advantage of the PH-FM estimator is that with it Hansen (1991) has developed tests of parameter stability of cointegrating relations. All the tests to be considered treat the break point, $\langle\pi T\rangle$, $\pi \in (0,1)$, as unknown.[18] $\langle\pi T\rangle$ represents the integer part of πT. In order to construct the tests, the data set must be trimmed to guarantee the test statistics do not diverge to infinity (almost surely). Andrews (1990) suggests $\pi \in [0.15,\ 0.85] \equiv \Pi$ and that is followed throughout this paper.

The null hypothesis of the first test is $H_o: \phi_1 = \phi_2$ and the alternative is $H_a: \phi_1 \neq \phi_2$, π unknown, where the subscripts denote either the first or second subsample. The test statistic is SupF = $\mathrm{Sup}_{\pi\in\Pi}$ F(π) where

$$F(\pi) = tr\left(\Xi_T(\pi)'\Sigma_T(\pi)^{-1}\Xi_T(\pi)\hat{\Omega}_{\eta\cdot\zeta}^{-1}\right)$$

$\Xi_T(\pi) = \sum_{t=1}^{\langle\pi T\rangle}\hat{\xi}_t$, $\hat{\Omega}_{\eta\cdot\zeta} = \hat{\Omega}_{\eta\eta} - \hat{\Omega}_{\eta\zeta}\hat{\Omega}_{\zeta\zeta}^{-1}\hat{\Omega}_{\zeta\eta}$, $\Sigma_T = M_T(\pi) - M_T(\pi)M_T(\pi)^{-1}M_T(\pi)$, and $M_T(\pi) = \sum_{t=1}^{\langle\pi T\rangle} y_t^2$. This test is referred to as the SupF test.

The assumption of a martingale process for ϕ is maintained for the other two tests of parameter instability of ϕ considered here. For $\phi_t = \phi_{t-1} + \varepsilon_t$, $\varepsilon_t \sim IID(0,\ S\sigma_\varepsilon^2)$, known constant S measures the parameter constancy during

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[18] As noted by Andrews (1990), treating the break point as known prior to testing sequentially for parameter instability causes the size of the tests for parameter constancy to become distorted because the limiting distribution of the test is altered.

date t. The $H_o: S = 0$ and $H_a: 0 < S$.[19] The F-test which is available under this null is given by $T^{-1}\sum_{t=1}^{<\pi T>} F(\pi)$ and is known as the Mean F test. With the martingale null a Lagrange multiplier like test exists as

$$Lc = tr\left(M_t(1)^{-1}\sum_{t=1}^{T}\hat{\xi}_t\hat{\Omega}_{\eta}^{-1}{}_{\zeta}\hat{\xi}_t\right).$$

An advantage of the Lc test is it requires no choice for the trim. Hansen (1991) provides details and asymptotic critical values for the three tests.[20]

A test for structural stability of the second round estimation of θ in (14) can be conducted with statistics developed by Ghysels and Hall (1990) and Andrews (1990). From the likelihood ratio test of Eichenbaum, Hansen, and Singleton (1988), Ghysels and Hall (1990) develop a statistic to test for parameter instability at a known fixed break point. The test statistic is

$$LR_T = J_{12}(\tilde{\theta}) - J_1(\hat{\theta})$$

where LR_T is asymptotically distributed $\chi^2(k-1)$, k-1 is the number of orthogonality conditions to be tested, $J_{12}(\tilde{\theta}) = J_1(\tilde{\theta}) + J_2(\tilde{\theta})$ is the sum of the quadratic GMM objective functions of (14) for the two subsamples evaluated at $\tilde{\theta}$, and $J_1(\hat{\theta})$ is the objective function of the GMM problem of (14) for the first subsample; see Ghysels and Hall (1990) for details.

Andrews (1990) allows for an unknown break point. The test statistic is a Wald statistic of the form $\mathrm{Sup}_{\pi\in\Pi} W_T(\pi)$ where

$$W_T(\pi) = T\left[\hat{\theta}_1(\pi) - \hat{\theta}_2(\pi)\right]\left[\pi^{-1}\hat{V}_{\theta GMM1}(\pi) + (1-\pi)^{-1}\hat{V}_{\theta GMM2}(\pi)\right]^{-1}\left[\hat{\theta}_1(\pi) - \hat{\theta}_2(\pi)\right],$$

where $\hat{V}_{\theta GMMi}$ is the estimate of the variance of the GMM estimator of θ for the i = 1, 2 subsample, for the $H_o: \theta_1 = \theta_2$ and $H_a: \theta_1 \neq \theta_2$. To implement the test, the trim is set by $\pi \in [0.15, 0.85]$; see Andrews (1990) for details and asymptotic critical values.

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[19] This choice for S means the hazard of parameter non-constancy is constant for the sample period.

[20] Gregory and Nason (1991) report small sample properties of the Mean F, SupF, and Lc tests in a variety of economic environments.

The second round estimates of ϕ and θ can be used to perform tests of the PIH with changing tastes. In all these tests, the uncertainty surrounding the estimates of ϕ and θ are ignored. The variance bounds inequality of (11) is computed in two steps. First, the variance of q_t is computed

$$\hat{\sigma}_q^2 = (T-1)^{-1}\sum_{t=1}^{T}\hat{q}_t^2,$$

where $\hat{q}_t = c_t - \hat{\phi}(1 - \hat{\theta})y_t - \hat{\theta}c_{t-1}$.

The second step is to estimate σ_e^2. To accomplish this, an ARI(p,1) for labor income is estimated in order to calculate $\rho([1+r]^{-1})$ and σ_e^2

$$(1-L)y_t = \mu + \sum_{i=1}^{p}\rho_i(1-L)y_{t-i} + e_t. \tag{16}$$

The variance of σ_e^2 is estimated with

$$\hat{\sigma}_e^2 = (T-2-p)^{-1}\sum_{t=1}^{T}\hat{e}_t^2,$$

where $\hat{e}_t$ is the date t OLS residual of (16). From (11), the variance bounds test statistic becomes

$$H_o:\ 0 \le \hat{\Psi}^2\hat{\rho}([1+r]^{-1})^{-2}\hat{\sigma}_e^2 - \hat{\sigma}_q^2, \tag{17}$$

where $\hat{\Psi}^2 \equiv \hat{\Gamma}^2(1 + \hat{\theta}^2)$. The standard errors of $\hat{\sigma}_q^2$, $\hat{\rho}([1+r]^{-1})^{-2}$, $\hat{\sigma}_e^2$, and the test statistics are computed numerically; for details see West (1988a).

Another test of the PIH with changing tastes can be constructed by regressing $\hat{q}_{t+1}$ on the elements of H_{t-1}, $H_{t-1} = ((1-L)y_{t-1}, \ldots, \hat{q}_{t-1}, \ldots)$. A simple means to implement this is the OLS regression

$$\hat{q}_{t+1} = \mu + \sum_{i=1}^{k}\psi_{yi}(1-L)y_{t-i} + \sum_{i=1}^{k}\psi_{qi}\hat{q}_{t-i} + \lambda_t. \tag{18}$$

The hypotheses tests are

$$H_{oy}:\ \psi_{yi} = 0, \quad i = 1, \ldots, k,$$

$$H_{oq}:\ \psi_{qi} = 0, \quad i = 1, \ldots, k,$$

and

$$H_{oyq}:\ \psi_{yi} = \psi_{qi} = 0, \quad i = 1, \ldots, k.$$

Wald tests are used to evaluate H_{oy} and H_{oq} and F-tests are used for H_{oyq}.

IV. Empirical Results

The data used for estimation has two sources. Time series for expenditures on services, nondurable goods, durable goods, and disposable income, yd_t, are from the National Income and Product Accounts (NIPA). The time series for aggregate labor income is from the MPS model data bank; see Brayton and Mauskopf (1985).[21] All series are quarterly, seasonally adjusted, per capita, and in 1982 dollars. The sample period is 1955.1 through 1987.4.

Two different definitions of per capita consumption are used in the empirical work. The first definition sets consumption equal to per capita expenditures on nondurable goods and services, c_{1t}. The other definition of consumption, c_{2t}, is per capita expenditures on nondurable goods, services, and durable goods and is referred to as total consumption expenditures.[22] The deflator used to convert per capita labor income, y_{1t} and y_{2t}, to 1982 dollars depends on the definition of per capita consumption.

To begin the empirical analysis, figure 1 contains time plots of c_{1t}/y_{1t}, c_{2t}/y_{2t} and for comparison c_{1t}/yd_t, and c_{2t}/yd_t. Not surprisingly given the results in Campbell (1987), c_{1t}/yd_t and c_{2t}/yd_t appear from casual inspection of figure 1 to be stationary time series. The same is true for c_{1t}/y_{1t}; however, this ratio exhibits much more variability than either c_{1t}/yd_t or c_{2t}/yd_t. A trend break appears to occur in c_{2t}/y_{2t} during either 1970 or 1974.

Table 1 contains the results of various univariate unit root tests. One means for testing for the presence of a unit root in a time series is to

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[21] The author wishes to thank Dawn Rehm for making the MPS data available.

[22] The PIH with changing tastes is tested with two different definitions of consumption. A reason for conducting inference in this fashion is to avoid some difficult empirical issues. The empirical tests of Eichenbaum and Hansen (1990) of the substitutability of durable goods for nondurable goods, when the preferences of the representative agent are quadratic, suggest these goods are perfect substitutes. However, aggregating durable and nondurable goods implies their relative price is fixed which, as Eichenbaum and Hansen (1990) note, is at odds with the data.

estimate the ADF regression

$$(1-L)x_t = \mu_1 + \mu_2 t + \delta x_{t-1} + \sum_{i=1}^{p}\nu_i(1-L)x_{t-i} + w_t$$

where p equals six in order to correct for the serial correlation in w_t.[23] The left most column of table 1 is the t-ratio, t_δ, of the estimate of δ. For neither c_{1t}, c_{2t}, y_{1t}, nor y_{2t}, is t_δ less than the 5 percent asymptotic critical level of -3.41.

The other unit root tests reported in table 2 have been studied by Phillips (1987) and Perron (1988). The Phillips-Perron regression is

$$(1-L)x_t = \mu_1 + \mu_2 t + \delta x_{t-1} + w_t.$$

The test statistics of the Phillips-Perron regression reported in table 2 have been corrected for serial correlation by a sixth order Newey-West (1987) correction.[24] The t-ratios, $Z(\delta)$ of the Phillips-Perron regression are listed in the second column (moving from right to left). The columns headed by $Z(\Phi_3)$ and $Z(\Phi_2)$ are the F-tests of the hypotheses $\mu_1 = \mu_2 = \delta = 0$ and $\mu_1 = \delta = 0$, respectively; see Perron (1988). As for the t-ratio from the ADF regression, none of these tests are able to reject the hypotheses of a unit root in either c_{1t}, c_{2t}, y_{1t}, or y_{2t}.

Given the evidence against trend stationarity in c_{1t}, c_{2t}, y_{1t}, and y_{2t}, several summary statistics for the first differences of c_{1t}, c_{2t}, y_{1t}, and y_{2t} are reported in table 2. Notably, the first-order correlation coefficient of $(1-L)c_{1t}$ and $(1-L)c_{1t-1}$ is 0.24 and slightly more than two standard deviations away from zero. Further, the contemporaneous correlation between $(1-L)c_{1t}$ and $(1-L)y_{1t}$ is nearly 0.6 and statistically significant while the correlation coefficient of $(1-L)c_{1t}$ and $(1-L)y_{1t-1}$ is less an 0.25 with a t-ratio slightly

[23] The results were little changed when either two or four lags of the first difference of the dependent variable were used.

[24] The Phillips-Perron test statistics were calculated with second and fourth order Newey-West corrections. The choice of the lag length for the Newey-West correction did not affect the test statistics greatly.

more than two.[25] Similar results are found for $(1-L)c_{2t}$ and $(1-L)y_{2t}$.

First round estimates of ϕ and θ appear in table 3. For c_{1t} and y_{1t} and then c_{2t} and y_{2t}, the OLS regression of (13) is run and the three tests of the null of consumption and labor income not being cointegrated are constructed. All of these test statistics provide little if any evidence against the null of no cointegration. Further, the PH-FM and the GMM estimates of ϕ are close to the OLS estimate of ϕ. For $\hat{\phi}_{GMM}$, the instrument matrix G is chosen to have the tth row $[1\ y_{t-2} \cdots y_{t-4}]$.

GMM estimation of (14) is performed with the instrument matrix

$$[1\ (1-L)y_{\phi t-2} - \alpha' z_{t-3} \cdots (1-L)y_{\phi t-4} - \alpha' z_{t-5}]$$

where the estimate of ϕ is the OLS estimate 1.083. Given stationarity of the dependent variable and the independent variable $(1-L)y_{1t}-\alpha' z_{1t-1}$, the over-identifying conditions cannot be rejected as the Hansen (1982) test statistic can only be rejected at the 0.2 level. At the bottom of table 3, the results for c_{2t} and y_{2t} are not qualitatively different.

Table 4 contains the second round estimates of ϕ and θ. For c^*_{1t} and y_{1t}, the most striking difference between the first and second round estimates of ϕ are the cointegration test statistics t_δ, Z_δ, and Z_t all reject the null hypothesis of no cointegration at least at the ten percent level in the second round. To generate c^*_{1t}, the first round estimate of θ of 0.956 is used. The results in the bottom half of table 4 tell the same story with the null of no cointegration rejected at the five percent critical levels for all the tests considered. To create c^*_{2t}, $\hat{\theta}$ equals 0.940.

Table 4 reports the SupF, Mean F, and Lc tests discussed in section III.

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[25] Christiano, Eichenbaum, and Marshall (1991) point out that temporal aggregation can be the source of the serial correlation in the first difference of consumption and the contemporaneous correlation between the first difference of consumption and the first difference of labor income.

None of the tests are able to reject any of the null hypotheses of parameter constancy of ϕ for either c_{1t} or c_{2t} as the dependent variable. This is true at the five and ten percent significance levels.[26]

To generate the second round GMM estimate of θ from (14) for c_{1t} and y_{1t}, the same instrument matrix is used as for the first round except the OLS estimate of ϕ equals 1.171. The second round GMM estimate of θ equals 0.974 and is significantly different from one. The three orthogonality conditions cannot be rejected at anything less than a 19 percent significance level. Roughly similar results appear in the bottom half of table 4 for c_{2t} and y_{2t}.

Tests of parameter constancy of θ are reported in table 4. The test statistic labeled LR is the Ghysels and Hall (1990) likelihood ratio test of parameter drift discussed in section III. This test requires a date be chosen to split the sample into two subsamples. A straightforward split is to break the sample into two equal halves. The split date is 1970.4. and roughly corresponds to the date when the trend of the ratio c_{2t}/y_{2t} *appears* to break in figure 1.[27] With three degrees-of-freedom (three over-identifying conditions are tested), the LR test statistic indicates rejection of the null hypothesis at the 10 percent level when $(1-L)c_{2t}$ is the dependent variable. For $(1-L)c_{1t}$, the null hypothesis is rejected only at the 12.6 percent level.

The other test of parameter constancy is the $\text{Sup}_{\pi\in\Pi}\ W_T(\pi)$ test of Andrews (1990) outlined in section III. This test statistic is denoted by W(1) because the test statistic has one degree-of-freedom (one parameter is estimated).

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[26] The critical levels Hansen (1991) reports are for the case of an intercept in the cointegration equation. However, equation (15) involves no intercept which means no nuisance parameter is estimated along with the parameter of cointegration. In this case, the five and ten percent critical levels reported at the bottom of table 4 for the SupF, Mean F, and Lc tests are conservative.

[27] The reason the Ghysels-Hall LR test is correctly sized, even though the break point is chosen prior to testing, is that only one date and not a sequence of dates is tested for parameter constancy.

Results similar to those for the Ghysels-Hall LR test are reported for W(1). Parameter constancy for the second round estimate of θ can be rejected only at significance levels greater than the ten percent level when $(1-L)c_{1t}$ is the dependent variable. When $(1-L)c_{2t}$ is the dependent variable, parameter stability of θ is rejected at less than the five percent critical level. In the lower half of table 4, the date corresponding to $<\pi T>$ of W(1) is 1982.2.

The tests of the PIH with changing tastes require construction of $\hat{q}_{1t}$ and $\hat{q}_{2t}$. $\hat{q}_{1t}$ is built with the second round estimates of $\hat{\phi} = 1.171$ and $\hat{\theta} = 0.974$ and for $\hat{q}_{2t}$ $\hat{\phi} = 1.367$ and $\hat{\theta} = 0.979$ have been used. To implement the variance bounds inequality test (17), the 'best' autoregressive model for $(1-L)y_{1t}$ and $(1-L)y_{2t}$, equation (16) is estimated by OLS for p = 0, 1, ..., 5 and the Akaike Information Criterion (AIC) is calculated; see Granger and Newbold (1986). For $(1-L)y_{1t}$ and $(1-L)y_{2t}$, the AIC picked p = 1 as the best model.

The estimates of the variance bounds inequality of table 5 were performed with quarterly real interest rates of 0.005 and 0.010. West (1988b) reports the choice of the real interest rate has little effect on the estimate of the variance bounds test statistic for the standard PIH. For the variance bounds test statistic (17), this is not true because the parameter Ψ^2 is sensitive to the choice of r. Although for either choice of r and for $\hat{q}_{1t}$ and $(1-L)y_{1t}$ the point estimates of the test statistic of (17) do not violate the variance bounds inequality and the t-ratios are greater than 1.96, the t-ratio is about two times larger for r equal to 0.005 than for r equal to 0.010.

For q_{2t} and $(1-L)y_{2t}$, the variance bounds inequality is satisfied for r equal to either 0.010 or 0.005. As before the point estimate of the test statistic of (17) is not violated for either choice of r, but now neither t-ratio is less than 4.1. Nonetheless, the evidence of parameter instability in θ suggests these estimates be read cautiously.

The results of estimating equation (18) by OLS appear in table 6. Table 6 contains the estimated sum of the coefficients ψ_{yi} and ψ_{qi}, their (numerically calculated) standard errors, Wald tests of H_{oy} and H_{oq}, and F-tests of H_{oyq}. In summary, the *hypothesis that variables in* H_{t-1} *provide no information for* predicting $\hat{q}_{t+1}$ can be rejected. Lags of $\hat{q}_1$ begin to matter for $\hat{q}_{1t+1}$ at k = 2. This occurs at k = 4 for the Wald test of H_{oy}; however, the t-ratio of the sum of $\hat{\psi}_{yi}$ is only slightly greater than one at k = 4 and not much greater than two at k = 5. The results are not qualitatively different for $\hat{q}_{2t+1}$, but lags of $\hat{q}_2$ and $(1-L)y_2$ begin to matter for $\hat{q}_{2t+1}$ much earlier. Once again, the regression results for $\hat{q}_2$ should be interpreted cautiously because of parameter instability in θ.

V. Conclusion

The purpose of this paper is to present a version of the permanent income model in which the bliss point of the agent is stochastic. Given the parametric form of the stochastic bliss point, a result of the PIH with changing tastes is that a linear combination of current consumption, current labor income, and lagged consumption form a stationary relationship. Further, if labor income is stationary in first differences, the PIH with changing tastes predicts consumption and labor income are cointegrated.

In order to conduct tests of the PIH with changing tastes the parameter of cointegration must be estimated as well as a parameter capturing the short run persistence in consumption. However, a bias in the tests for cointegration exists because of serial correlation in consumption. A transformation exists which rids this regression of the serial correlation in the residuals, but this transformation requires a two round estimation procedure. It is possible to recover the parameter of cointegration and the persistence parameter from this two round procedure. The results of this two round procedure reject the null hypothesis that consumption and labor income are not cointegrated for either of the definitions of consumption used in this paper.

Tests of the parameter constancy of the parameter of the cointegration relation and the persistence parameter are conducted. The tests suggest parameter constancy for the parameter of cointegration irrespective of the definition of consumption. For the persistence parameter, parameter constancy is not rejected for consumption defined as expenditures on services and nondurable goods and is rejected for total consumption.

The results of the tests of the PIH with changing tastes are mixed at best. The PIH with changing tastes identifies explicitly the permanent and temporary shocks of the bivariate stochastic process of consumption and labor income. The identification scheme of the PIH with changing tastes is the long run equilibrium (cointegration relation) that exists between consumption and labor income and the short run persistence in consumption. Tests imposing this identification scheme only go part of the way in resolving the excess smoothness quandary. These tests suggest time variation in real interest rates have implications for the observed smoothness of consumption. Nonetheless, consumption appears to be more serially correlated than the PIH with changing tastes is capable of supporting.

It appears that the PIH with changing tastes begins to explain or at least provides a partial explanation for some of the empirical anomalies often reported for aggregate consumption and labor income. This might recommend the PIH with changing tastes as an alternative to the standard PIH. Nonetheless, this is a misreading of the model and empirical results that are found in this paper. The successes and failures of the PIH with changing tastes give only some direction in which to improve the dynamic model of aggregate household behavior. In the context of the results of this paper, the problems remaining are to integrate time-varying real interest rates into the model as well as an explicit decision for durable goods expenditures. Attempts to model and test the implications of time-varying real interest rates and the decision to purchase durables goods are left for future research.

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Appendix

This appendix constructs the variance bounds inequality (12). To begin, define $x_t = c_t - \theta c_{t-1} - \theta(1-\phi)y_t$ and

$$x_t^D = c_t^D - \theta c_{t-1}^D - \theta(1-\phi)y_t^D = \sum_{j=0}^{\infty}(1+r)^{-j}\Big[c_t - \theta c_{t-1} - \theta(1-\phi)y_t\Big].$$

This implies $x_t^D = x_t + (1+r)^{-1}x_{t+1}^D$. In turn,

$$x_t^I = x_t + (1+r)^{-1}E_t x_{t+1}^D \tag{A1}$$

and

$$x_t^H = x_t + (1+r)^{-1}E_t^H x_{t+1}^D. \tag{A2}$$

The other definitions needed are $f_{t+1} = x_{t+1}^I - E_t x_{t+1}^D = x_{t+1}^I - E_t x_{t+1}^I$ and $g_{t+1} = x_{t+1}^H - E_t^H x_{t+1}^D = x_{t+1}^H - E_t^H x_{t+1}^H$. By combining the last set of definitions with (A1) and (A2), the results are $x_t^I = x_t + (1+r)^{-1}[x_{t+1}^I - f_{t+1}]$ and $x_t^H = x_t + (1+r)^{-1}[x_{t+1}^H - g_{t+1}]$. A forward recursion produces

$$x_t^I = \sum_{j=0}^{\infty}(1+r)^{-j}x_{t+j} - \sum_{j=1}^{\infty}(1+r)^{-j}f_{t+j} \tag{A3}$$

and

$$x_t^H = \sum_{j=0}^{\infty}(1+r)^{-j}x_{t+j} - \sum_{j=1}^{\infty}(1+r)^{-j}g_{t+j}. \tag{A4}$$

From proposition 1 of West (1988a),

$$E\Big[x_t^H - E\{x_t^H/H_{t-1}\}\Big]^2 \ge E\Big[x_t^I - E\{x_t^I/I_{t-1}\}\Big]^2.$$

Hence, (A3) and (A4) result in

$$\sum_{j=1}^{\infty}(1+r)^{-2j}Eg_t^2 \ge \sum_{j=1}^{\infty}(1+r)^{-2j}Ef_t^2$$

and it follows $Eg_t^2 \ge Ef_t^2$. Finally, note

$$Ef_t^2 = E\Big[q_t - E\{q_t/H_{t-1}\}\Big]^2 = \sigma_q^2 \tag{A5}$$

and

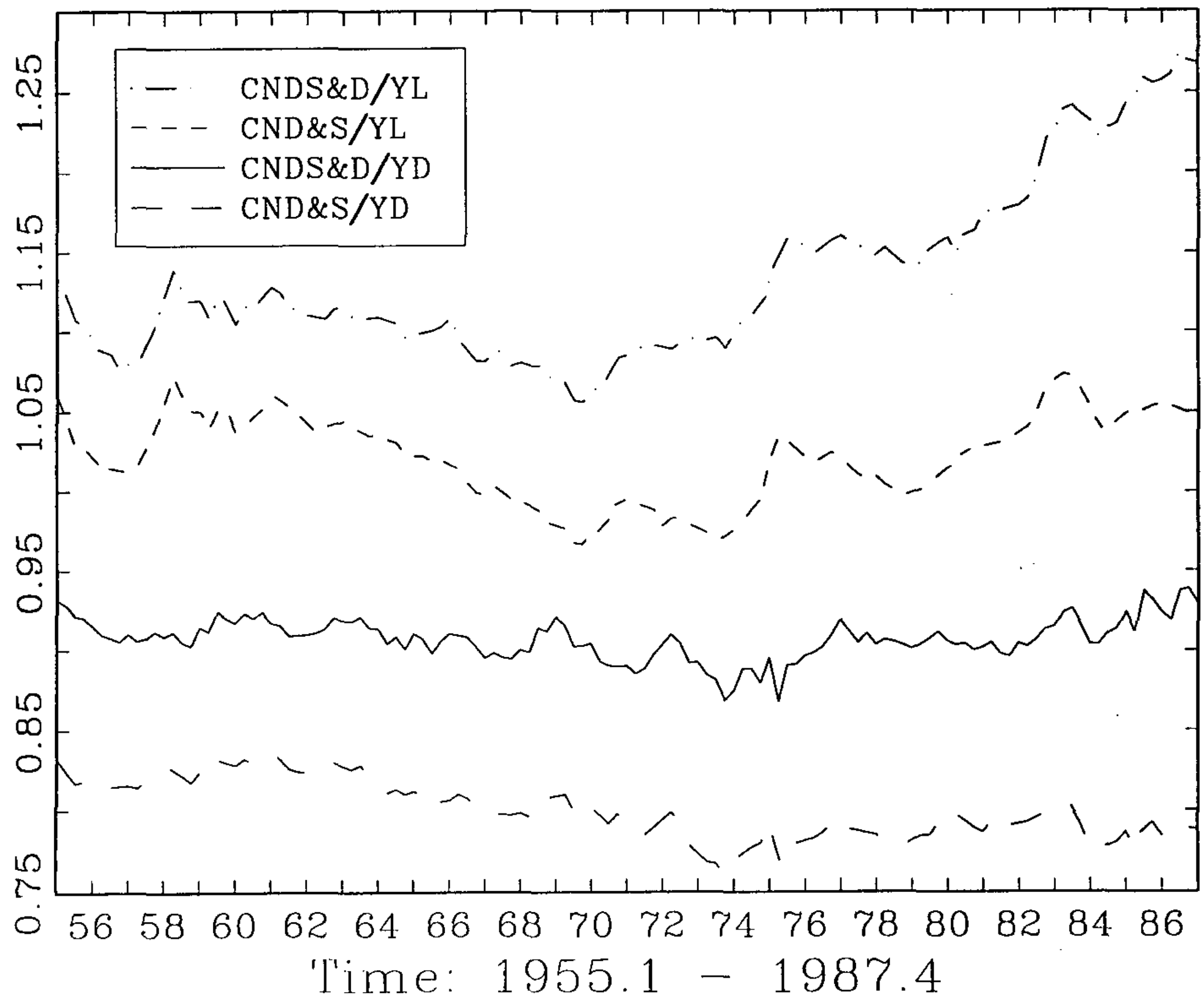
$$Eg_t^2 = \Gamma^2\Big[\Big[y_t^H - E\{y_t^H/H_{t-1}\}\Big] - \theta\Big[y_{t-1}^H - E\{y_{t-1}^H/H_{t-2}\}\Big]\Big]^2$$

where the formula of Hansen and Sargent (1980), (16), and e_t being serially uncorrelated yield

$$Eg_t^2 = \Gamma^2(1+\theta^2)\Big[\rho([1+r]^{-1})^{-1}e_t\Big]^2. \tag{A6}$$

The variance bounds inequality of (11) follows from (A5) and (A6).

FIGURE 1: C/Y



CNDS&D/YL
CND&S/YL
CNDS&D/YD
CND&S/YD
0.75
0.85
0.95
1.05
1.15
1.25
56 58 60 62 64 66 68 70 72 74 76 78 80 82 84 86
Time: 1955.1 – 1987.4

Table 1.

Tests For Unit Roots: Estimated t-ratios and F-tests
Estimation Period: 1955.1 to 1987.4

Series	t_δ^1	$Z(\delta)^2$	$Z(\Phi_3)^2$	$Z(\Phi_2)^2$
c_1	-2.926	-2.412	21.232	3.353
c_2	-3.007	-2.437	13.338	3.818
y_1	-2.019	-1.803	4.911	1.693
y_2	-2.285	-1.993	6.282	1.997

Definitions:

c_1 = expenditures on nondurables and services in constant 1982 dollars (NIPA),

c_2 = expenditures on nondurables, services, and durables in constant 1982 dollars (NIPA),

y_1= labour income deflated by the price deflator for expenditures on nondurables and services into constant 1982 dollars (MPS),

y_2= labour income deflated by the price deflator for expenditures on nondurables, services, and durables into constant 1982 dollars (MPS).

---

1 The augmented Dickey-Fuller regression is

$$(1-L)x_t = \mu_1 + \mu_2 + \delta x_{t-1} + \sum_{i=1}^{6}(1-L)\upsilon_i x_{t-i} + w_t.$$

t_δ is the t-ratio of δ. The asymptotic five percent critical value is -3.41 and at ten percent it is -3.12; see Fuller (1976).

2 The Phillips-Perron regression is

$$(1-L)x_t = \mu_1 + \mu_2 t + \delta x_{t-1} + w_t$$

with a Newey-West correction of lag length 6 to obtain the asymptotic standard errors and F-statistics. $Z(\rho)$ is the t-ratio of the estimate of ρ. $Z(\Phi_3)$ is the F-test of $\mu_1 = \mu_2 = \delta = 0$. $Z(\Phi_2)$ is the F-test of $\mu_1 = \delta = 0$. The asymptotic five percent critical value is -3.41 and at ten percent it is -3.12 for the t-ratios $Z(\rho)$. The five and ten percent critical values for the F-test $Z(\Phi_3)$ are 6.25 and 5.34, respectively. For $Z(\Phi_2)$, the relevant critical values are 4.68 and 4.03 at the five and ten percent levels; see Perron (1988).

Table 2.

Means, Standard Deviations, and Correlations
Estimation Period: 1955.1 - 1987.4

	Mean	Standard Deviation
$(1 - L)c_{1t}$	31.374	38.329
	(4.238)[1]	(3.014)
$(1 - L)y_{1t}$	27.909	53.137
	(8.090)	(2.908)
$(1 - L)c_{2t}$	39.372	55.308
	(7.098)	(4.558)
$(1 - L)y_{2t}$	30.402	51.365
	(7.740)	(3.126)

Correlation Matrix

$(1 - L)c_{1t}$	1.000	0.244	0.592	0.234
		(0.119)	(0.090)	(0.119)
$(1 - L)c_{1t-1}$		1.000	0.412	0.602
			(0.093)	(0.091)
$(1 - L)y_{1t}$			1.000	0.522
				(0.086)
$(1 - L)y_{1t-1}$				1.000

Correlation Matrix

$(1 - L)c_{2t}$	1.000	0.232	0.579	0.248
		(0.115)	(0.083)	(0.129)
$(1 - L)c_{2t-1}$		1.000	0.515	0.615
			(0.097)	(0.082)
$(1 - L)y_{2t}$			1.000	0.545
				(0.091)
$(1 - L)y_{2t-1}$				1.000

---

[1] All numbers in parentheses are asymptotic heteroskedastic, autocorrelation consistent standard errors. To obtain them, a sixth order Newey-West correction was used.

Table 3.
First Round Estimates of Cointegrating Vector, Tests of Cointegration and Error Correction Models
Estimation Period: 1955.1 to 1987.4

Consumption Expenditures on Nondurables and Services[1]

OLS: $c_{1t} = \underset{(0.020)}{1.083} y_{1t}, R^2 = 0.958$

$t_\delta^2 = -1.393$ with two lags, -1.241 with six lags, $Z_\delta^3 = -3.601$, and $Z_t^4 = -1.215$.

PH-FM: $c_{1t} = \underset{(0.064)}{1.090} y_{1t}$

GMM: $c_{1t} = \underset{(0.045)^5}{1.091} y_{1t},\ J(3) = \underset{(0.198)^6}{4.668}$

GMM: $(1 - L)c_{1t} = \underset{(0.006)}{0.044} [(1 - L)y_{1\phi t} - \alpha' z_{1t-1}],\ J(3) = \underset{(0.208)}{4.552},\ \theta = \underset{(0.006)}{0.956}$

Total Consumption Expenditures[1]

OLS: $c_{2t} = \underset{(0.019)}{1.271} y_{2t}, R^2 = 0.971$

$t_\delta^2 = -1.493$ with two lags, -1.061 with six lags, $Z_\delta^3 = -4.149$, and $Z_t^4 = -1.393$.

PH-FM: $c_{2t} = \underset{(0.063)}{1.282} y_{2t}$

GMM: $c_{2t} = \underset{(0.047)^5}{1.279} y_{2t},\ J(3) = \underset{(0.423)^6}{2.801}$

GMM: $(1 - L)c_{2t} = \underset{(0.012)}{0.060} [(1 - L)y_{2\phi t} - \alpha' z_{2t-1}],\ J(3) = \underset{(0.155)}{5.237},\ \theta = \underset{(0.012)}{0.940}$

---

1 All variables have been demeaned prior to estimation.

2 The augmented Dickey-Fuller regression is

$$(1 - L)\hat{\eta}_t = \delta\hat{\eta}_{t-1} + \sum_{i=1}^{p} \upsilon_i (1 - L)\hat{\eta}_{t-i} + w_t,$$

where $\eta_t = c_{it} - \hat{\phi} y_{it}$, $i = 1, 2$. As found in table IIb of Phillips and Ouliaris (1990) the critical values at the five and ten percent significance levels for t_δ are -3.768 and -3.449, respectively.

3 Table Ib of Phillips and Ouliaris (1990) report critical values for Z_δ at the five and ten percent significance level of -21.483 and -18.178, respectively. A sixth order Newey-West correction was used.

4 Table IIB *of Phillips and Ouliaris report critical values for* Z_t *at the five and ten percent significance level* of -3.768 and 3.449, respectively.

5 The number in parenthesis is the asymptotic heteroskedastic, autocorrelation consistent standard error. A 12th order Newey-West correction was used.

6 The asymptotic confidence level is in parenthesis.

Table 4.

Second Round Estimates of Cointegrating Vector, Tests of Cointegration and Error Correction Models Estimation Period: 1955.1 to 1987.4

Consumption Expenditures on Nondurables and Services[1]

OLS: $c^{*1}_{1t} = \underset{(0.063)}{1.171} y_{1t}, R^2 = 0.728$

t^2_δ = -4.016 with two lags, -3.195 with six lags, $Z^3_\delta = -95.217$, and $Z^4_t = -8.373$.

PH-FM: $c^{*1}_{1t} = \underset{(0.072)}{1.189} y_{1t}$, SupF = $\underset{(1982.4)^5}{10.522}$, Mean F = 1.792, Lc = 0.339

GMM: $c^{*1}_{1t} = \underset{(0.065)^6}{1.175} y_{1t}$, $J(3) = \underset{(0.439)^7}{2.706}$

GMM: $(1 - L)c_{1t} = \underset{(0.004)}{0.026} [(1 - L)y_{1\phi t} - \alpha' z_{1t-1}]$, $J(3) = \underset{(0.192)}{4.741}$, $\theta = \underset{(0.004)}{0.974}$, $LR(3) = \underset{(0.126)}{5.713}$, $W(1)^8 = 6.564$

Total Consumption Expenditures[1]

OLS: $c^{*1}_{2t} = \underset{(0.019)}{1.367} y_{2t}, R^2 = 0.738$

t^2_δ = -4.012 with two lags, -3.406 with six lags, $Z^3_\delta = -97.886$, and $Z^4_t = -8.406$

PH-FM: $c^{*1}_{2t} = \underset{(0.073)}{1.367} y_{2t}$, SupF = $\underset{(1982.3)^5}{9.961}$, Mean F = 1.967, Lc = 0.375

GMM: $c^{*1}_{2t} = \underset{(0.074)^6}{1.344} y_{2t}$, $J(3) = \underset{(0.494)^7}{2.398}$

GMM: $(1 - L)c_{2t} = \underset{(0.005)}{0.021} [(1 - L)y_{2\phi t} - \alpha' z_{2t-1}]$, $J(3) = \underset{(0.229)}{4.316}$, $\theta = \underset{(0.005)}{0.979}$, $LR(3) = \underset{(0.105)}{6.140}$, $W(1)^8 = 8.887$

---

1 $c^{*1}_{it} \equiv c_{it} + \theta_i(1-\theta_i)^{-1}(1-L)c_{it}$, $i = 1, 2$, $\theta_1 = 0.956$, and $\theta_2 = 0.940$.

2 See footnote 2 table 3.

3 See footnote 3 table 3.

4 See footnote 4 table 3.

5 The five percent critical values found in Hansen (1991) for the SupF, Mean F, and Lc statistics are 12.0, 4.17, and 0.582, respectively. The number in parenthesis is the date corresponding to the SupF statistic.

6 See footnote 5 table 3.

7 See footnote 6 table 3.

8 The five and ten percent critical values of the SupF Wald Test with one degree-of-freedom given in Andrews (1990) are 8.7 and 7.2 percent, respectively.

Table 5.

Variance Bounds Test
Estimation Period: 1955.1 to 1987.4

Consumption Expenditures on Nondurables and Services

	μ	ρ	σ_e^2	$\rho[(1+r)^{-1}]^{-1}$	σ_H^2	Ψ^2	$\Psi^2\sigma_H^2$	$\sigma_{q_1}^2$	$\Psi^2\sigma_H^2 - \sigma_{q_1}^2$
r = 0.005:	13.351	0.523	2410.160	2.084	10462.766	2.064	21597.919	1221.892[2]	20376.027
	(4.228)[1]	(0.049)	(555.428)	(0.213)	(3222.611)	-	(4630.103)	(398.555)	(4647.225)
r = 0.010:	13.351	0.523	2410.160	2.072	10351.418	0.689	7133.802	1221.892	5911.909
	(4.228)	(0.049)	(555.428)	(0.210)	(3173.958)	-	(2634.886)	(398.555)	(2664.858)

Total Consumption Expenditures

	μ	ρ	σ_e^2	$\rho[(1+r)^{-1}]^{-1}$	σ_H^2	Ψ^2	$\Psi^2\sigma_H^2$	$\sigma_{q_2}^2$	$\Psi^2\sigma_H^2 - \sigma_{q_2}^2$
r = 0.005:	14.016	0.551	2004.967	2.212	9808.544	6.898	67660.087	3263.803[3]	64396.284
	(3.928)	(0.046)	(506.592)	(0.222)	(3162.846)	-	(8306.960)	(651.379)	(8332.459)
r = 0.010	14.016	0.551	2004.967	2.199	9691.910	2.422	23472.164	3263.803	20208.361
	(3.928)	(0.046)	(506.592)	(0.218)	(3112.150)	-	(4843.200)	(651.379)	(4886.807)

1 A twelfth order Newey-West correction was used to generate the asymptotic heteroskedastic, autocorrelation consistent standard errors in parentheses.

2 $q_{1t} = c_{1t} - 0.974c_{1t-1} - 0.031y_{1t}$.

3 $q_{2t} = c_{2t} - 0.979c_{2t-1} - 0.029y_{2t}$.

Table 6
Tests of Unpredictibility of q_{t+1}
Estimation Period: 1955.1 to 1987.4

Consumption Expenditures on Nondurable Goods and Services[1]

k	$\sum_{i=1}^{k} \psi_{y_i}$	$\sum_{i=1}^{k} \psi_{q_i}$	$W_k(y)^2$	$W_k(q)$	$F(2k, T-2k-1)^3$
1	0.042 (0.060)[4]	0.116 (0.115)	0.488 (0.485)[5]	1.017 (0.313)	1.811 (0.168)
2	0.021 (0.071)	0.317 (0.119)	1.223 (0.542)	7.089 (0.029)	2.770 (0.030)
3	-0.055 (0.075)	0.417 (0.131)	3.842 (0.279)	11.577 (0.009)	2.397 (0.032)
4	-0.078 (0.074)	0.348 (0.183)	8.306 (0.081)	10.085 (0.039)	2.639 (0.011)
5	-0.153 (0.075)	0.482 (0.188)	17.400 (0.004)	17.213 (0.004)	2.247 (0.019)

Total Consumption Expenditures[6]

k	$\sum_{i=1}^{k} \psi_{y_i}$	$\sum_{i=1}^{k} \psi_{q_i}$	$W_k(y)^2$	$W_k(q)$	$F(2k, T-2k-1)^3$
1	-0.037 (0.087)[4]	0.240 (0.110)	0.178 (0.673)[5]	4.712 (0.030)	3.184 (0.045)
2	-0.152 (0.136)	0.523 (0.146)	3.473 (0.176)	12.759 (0.002)	3.961 (0.005)
3	-0.371 (0.114)	0.680 (0.107)	17.247 (0.001)	41.976 (0.000)	4.369 (0.000)
4	-0.410 (0.114)	0.639 (0.167)	21.393 (0.000)	21.936 (0.000)	3.734 (0.001)
5	-0.516 (0.162)	0.747 (0.203)	35.352 (0.000)	27.244 (0.000)	3.061 (0.002)

1 $q_{1t} = c_{1t} - 0.974c_{1t-1} - 0.031y_{1t}$.

2 The subscript indicates the degrees of freedom of the test statistic.

3 T is the number of observations, 132.

4 Newey-West (1987) asymptotic heteroskedastic, autocorrelation consistent standard errors are in parentheses. A twelth order correction was used.

5 The p-value in parentheses is in percent.

6 $q_{2t} = c_{2t} - 0.979c_{2t-1} - 0.029y_{2t}$.