JOURNAL OF APPLIED ECONOMETRICS, VOL. 11, 119–134 (1996)

A TIME SERIES ANALYSIS OF REAL WAGES, CONSUMPTION, AND ASSET RETURNS

THOMAS F. COOLEY

Department of Economics and Simon School of Business, University of Rochester, Rochester, NY 14627, USA

MASAO OGAKI

Department of Economics, 410 Arps Hall, Ohio State University, 1945N. High Street, Columbus, OH 43210, USA
Email: mogaki@ecolan.sbs.ohio-state.edu

SUMMARY

This paper re-examines whether the time series properties of aggregate consumption, real wages, and asset returns can be explained by a neoclassical model. Previous empirical rejections of the model have suggested that the optimal labour contract model might be appropriate for understanding the time series properties of the real wage rate and consumption. We show that an optimal contract model restricts the long-run relation of the real wage rate and consumption. We exploit this long-run restriction (cointegration restriction) for estimating and testing the model, using Ogaki and Park's (1989) cointegration approach. This long-run restriction involves a parameter that we call the long-run intertemporal elasticity of substitution (IES) for non-durable consumption but does not involve the IES for leisure. This allows us to estimate the long-run IES for non-durable consumption from a cointegrating regression. Tests for the null of cointegration do not reject our model. As a further analysis, our estimates of the long-run IES for non-durable consumption are used to estimate the discount factor and a coefficient of time-nonseparability using Hansen's (1982) Generalized Method of Moments. We form a specification test for our model *à la* Hausman (1978) from these two steps. This specification test does not reject our model.

1. INTRODUCTION

This paper re-examines whether the time series properties of aggregate consumption, real wages, and asset returns are consistent with a simple neoclassical representative agent economy. Previous empirical explorations of this issue have rejected the neoclassical model in large part because the marginal rate of substitution between consumption and leisure does not equal the real wage as is implied by the the first-order conditions of the model. In this paper we argue that an optimal labour contracting model is more appropriate for understanding the time series behaviour of real wages and consumption. We show that a version of the optimal contract model restricts the long-run relation between real wages and consumption. We exploit this long-run restriction (cointegration restriction) to estimate preference parameters and test the model.[1] First, we employ the cointegration approach suggested by Ogaki and Park (1989) to estimate the long-run intertemporal elasticity of substitution for non-durable consumption from a cointegrating regression. We test the model by testing for the cointegration restriction.

[1] Abowd and Card (1987) used micro data to compare a contracting model and a neoclassical labour supply model. Their focus was on earning changes and hours changes and very different from ours.

CCC 0883–7252/96/020119–16

Received June 1994
Revised April 1995

As a further analysis, we use this estimated preference parameter in the asset pricing equation implied by this economy to estimate the discount factor and a coefficient of time-non-separability using Hansen's (1982) Generalized Method of Moments (GMM). From this we are able to construct another specification test of the model.

Mankiw, Rotemberg, and Summers (1985, hereafter Mankiw *et al.*) subjected the Euler equations of an intertemporal labour supply model to a battery of tests and found no evidence to support it. Not only did their formal tests reject the model, but their point estimates of preference parameters implied a convex utility function. They concluded that the observed 'economic fluctuations do not easily admit of a neoclassical interpretation'.

Eichenbaum, Hansen, and Singleton (1988, hereafter Eichenbaum *et al.*) also used the Euler equation approach, but their point estimates of preference parameters were more reasonable. They attributed their different finding to two factors. First, they removed trends by taking growth rates of variables and taking ratios of variables while Mankiw *et al.* did not. Second, Eichenbaum *et al.* allowed time-nonseparability of preferences. Though their point estimates were reasonable, their formal test statistics typically rejected the model at the 1 per cent level when they tested both asset pricing equations and the first-order condition that equates the real wage with the marginal rate of substitution between leisure and consumption. When they removed the first-order condition and tested the asset pricing equations, their tests did not reject the model. However, the loss of precision of their estimates was substantial when the first-order condition was removed. Eichenbaum *et al.* interpreted their results as suggesting that the optimal labour contract model might be appropriate for understanding real wages.[2]

A given Pareto optimal allocation can be consistent with a wide variety of institutional arrangements. In optimal labour contract models (see, e.g. Azariadis, 1975, Rosen, 1985; Wright, 1988, labour income contains a component that provides workers with some degree of protection against business cycle fluctuations (also see Hall, 1980). This insurance component of labour income inserts a wedge between the marginal rate of substitution between leisure and consumption and wages. In their empirical work, Gomme and Greenwood (1995) showed that accounting for this component could help explain the observed pattern of fluctuations in income. These arguments combined with the findings of Eichenbaum *et al.* suggest that the imposition of the requirement that wages equal the marginal rate of substitution between consumption and leisure is too confining.

In the present paper, we use a restriction on the time series properties of real wages and consumption that is implied by optimal labour contracting to estimate preference parameters and test the model. In the optimal contract model, the first-order condition for real wages and consumption does not hold on a period-by-period basis. We will show, however, that a version of the optimal contract model implies that the real wage rate is equated with the marginal rate of substitution between consumption and labour in the long run. We exploit this long-run restriction for estimating and testing the model.

In contrast to the research cited above, the cointegration approach yields results that are supportive of the representative agent model. In the first step of our econometric procedure, we test the null hypothesis of cointegration and estimate the long-run IES for three measures of non-durable consumption. We do not reject the null of cointegration and obtain reasonable estimates. The long-run IES appears in the asset pricing equation derived from the representative consumer model. We use our estimated IES parameter in the asset pricing equation and apply

[2] Osano and Inoue (1991) use an approach similar to Eichenbaum *et al.* to test the overidentifying restrictions of Euler equations, using aggregate Japanese data. They also noted that there was much less evidence against the model when they removed the Euler equation associated with the equation of real wages and the marginal product of labour.

GMM to estimate the discount parameter and a coefficient of time-non-separability. We use both stock and nominal risk free returns. We form a specification test *à la* Hausman (1978) through these steps. This specification test does not reject the model.

The rest of the present paper is organized as follows. In Section 2, we present our model and derive the cointegration restriction. We describe our econometric procedures in Section 3. In Section 4, we explain the data. Section 5 contains our empirical results. Our conclusions are in Section 6.

2. THE ECONOMY

2.1. The Cointegration Restriction

We consider an economy populated by N households who have preferences defined over consumption and the flow of services from their leisure time. Each household maximizes

$$U = E_0\left[\sum_{t=0}^{\infty} \beta^t u(t)\right] \tag{1}$$

where E_t denotes the expectation conditioned on the information available at t. In order to develop intuition, let us first consider a simple intra-period utility function that is assumed to be time- and state-separable and separable in non-durable consumption, durable consumption, and leisure:

$$u(t) = \frac{C(t)^{1-\alpha} - 1}{1-\alpha} + v(l(t)) \tag{2}$$

where $v(\cdot)$ is a continuously differentiable concave function, $C(t)$ is non-durable consumption, and $l(t)$ is leisure.

For now, assume that real wages do not contain any insurance component. Then the usual first-order condition for a household that equates the real wage rate with the marginal rate of substitution between leisure and consumption is:

$$W(t) = \frac{v'(l(t))}{C(t)^{-\alpha}} \tag{3}$$

where $W(t)$ is the real wage rate. We assume that the stochastic process of leisure is (strictly) stationary in the equilibrium as in Eichenbaum *et al.* and that the random variables used to form the conditional expectations for stationary variables are stationary. Then an implication of the first-order condition is that $\ln(Wt) - \alpha \ln(C(t) = \ln(v'(l(t))$ is stationary. When we assume that the log of consumption is difference stationary, this implies that the log of the real wage rate and the log of consumption are cointegrated with a cointegrating vector $(1, -\alpha)'$. We exploit this cointegration restriction to identify the curvature parameter α from cointegrating regressions.

Given that the saving rate is stable in the long run in the USA (as Kuznets, 1946, found), it is natural to impose a restriction that the ratio of total consumption expenditure and labour income is stable at least when a consumer is rich enough. Because we assume that consumption and leisure are additively separable in intertemporal preferences, this restriction implies that α is equal to one when total consumption expenditure is used as $C(t)$ in our model. In our empirical work, we use a measure of non-durable consumption as $C(t)$, assuming that the other consumption goods (say, durable consumption goods) are additively separable from the measure of non-durable consumption good used in our analysis. For this reason, α can be different from one even when the saving rate is stationary for rich enough consumers.[3]

We now introduce time-non-separability of preferences. The intra-period utility function is assumed to be

$$u(t) = \frac{S(t)^{1-\alpha} - 1}{1-\alpha} + v(l(t), l(t-1), ..., l(t-k)) \tag{4}$$

where $S(t)$ is the service flow from non-durable consumption:

$$S(t) = C(t) + \lambda C(t-1) \tag{5}$$

This type of time non-separable specification of leisure has been used by many authors[4] and is useful because it can capture the fact that households may use leisure time in a household production technology to augment a stock of household capital (Kydland, 1984; Greenwood and Hercowitz, 1990; Benhabib, Rogerson, and Wright, 1991).

The time-non-separable specification for non-durable consumption is similar to that considered by Eichenbaum, Hansen, and Singleton (1988), Eichenbaum and Hansen (1990), Constantinides (1990), Heaton (1993, 1995), Allen (1992), and Braun, Constantinides, and Ferson (1993) among others, except that some of these authors considered more general form of time- non-separability for non-durable consumption than equation (5). We have habit formation for non-durable consumption when λ is negative and local substitutability or durability when λ is positive.[5] Note that the time- non-separability does not affect the IES in the long-run when $C(t)$ and $C(t-1)$ are equal.[6] We will refer to $1/\alpha$ as the long-run IES for non-durable consumption.

The usual first-order condition for a household that equates real wage rate with the marginal rate of substitution between leisure and consumption is now:

$$\begin{aligned} W(t) = \frac{\partial U/\partial l(t)}{\partial U/\partial C(t)} &= \frac{E_t\left[\sum_{\tau=0}^{K} \beta^\tau \partial u(t+\tau)/\partial l(t)\right]}{E_t[\partial u(t)/\partial C(t) + \partial u(t+1)/\partial C(t)]} \\ &= \frac{E_t\left[\sum_{\tau=0}^{K} \beta^\tau \partial v(t+\tau)/\partial l(t)\right]}{E_t[S(t)^{-\alpha} + \beta\lambda S(t+1)^{-\alpha}]} \end{aligned} \tag{6}$$

We assume that $\ln(C(t))$ is difference stationary in the equilibrium. Then

$$S(t+\tau)/C(\mathrm{t}) = C(t+\tau)/C(t) + \lambda C(t+\tau+1)/C(t) \tag{7}$$

[3] Because many economic models imply known cointegrating vectors when the log of the variables are taken and because an attractive feature of cointegration is that unknown parameters can be estimated without exogeneity assumptions, the fact that α is unknown in the model is important. For this reason, this point that α can be different from one is explained in details in the Appendix.

[4] See e.g., Kydland and Prescott (1982), Kennan (1988), Hotz, Kydland and Sedlacek (1988) and Eichenbaum, Hansen, and Singleton (1988).

[5] The time-non-separability for non-durable consumption allows us to separate the IES in the short run and the reciprocal of the RRA coefficient as Constantinides (1990) described, which could help explain the equity premium puzzle of Mehra and Prescott (1985). Ferson and Constantinides (1991) found evidence in favour of the asset pricing model with habit formation, using GMM.

[6] Alternatively, $C(t)$ grows at a constant rate in the long run.

is stationary for any τ. Combined with the first-order condition (6), this implies that

$$W(t)C(t)^{-\alpha} = \frac{E_t\left[\sum_{\tau=0}^{K} \beta^{\tau}\, \partial v(t+\tau)/\partial l(t)\right]}{E_t[\{S(t)/C(t)\}^{-\alpha} + \beta\lambda\{S(t+1)/C(t)\}^{-\alpha}]} \tag{8}$$

is stationary. Taking logs, $\ln(W(t)) - \alpha \ln(C(\mathrm{t}))$ s stationary as in the time-separable case we discussed.

2.2. Aggregation

We have derived the cointegration restriction for individual households. This restriction also holds for aggregated data under certain conditions when the restriction holds for each household i ($i = 1, \ldots, N$). Denoting household i's variable by a subscript i, let $W_a = (1/N)\Sigma_{i=1}^{N} W_i$ and $C_a = (1/N)\Sigma_{i=1}^{N} C_i$. The cointegration restriction implies $(1/N)\Sigma_{i=1}^{N} \log(W_i) - \alpha(1/N)\Sigma_{i=1}^{N} \log(C_i)$ is stationary but we observe $\log(W_a)$ and $\log(C_a)$. Thus a sufficient condition for aggregation of the cointegration restriction is that the difference between the average of log and the log of average over households of each of these variables is stationary.[7] This condition is empirically testable with panel data. The condition is satisfied for consumption under complete markets because the flow of services grows at the same rate for all consumers under the perfect risk sharing and because the ratio of service flow to consumption is stationary as we discussed.[8] The assumption of complete markets does not guarantee the aggregation condition for wage rates. This is because the difference in wage rates for households is mainly caused by the difference in abilities of workers. We assume that the distribution of abilities is such that the aggregation condition holds for the real wage rates.

In our empirical work, we estimate and test the first-order condition (6) through the cointegration restriction for aggregated real wages and consumption. We also estimate and test the standard asset pricing equation for the time-non-separable utility function

$$\frac{E_t[\beta\{S(t+1)^{-\alpha} + \beta\lambda S(t+2)^{-\alpha}\}R(t+1)]}{E_t[\{S(t)^{-\alpha} + \beta\lambda S(t+1)^{-\alpha}\}]} = 1 \tag{9}$$

for any gross asset return $R(t)$. Dividing both the numerator and denominator of equation (9) by $S(t)^{-\alpha}$, which is in the information set available at t, we obtain

$$\frac{E_t[\beta\{(S(t+1)/S(t))^{-\alpha} + \beta\lambda(S(t+2)/S(t))^{-\alpha}\}R(t+1)]}{E_t[1 + \beta\lambda\{S(t+1)/S(t)\}^{-\alpha}]} = 1 \tag{10}$$

This asset pricing formula is also satisfied by the aggregated service flow $S_a(t) \equiv (1/N)\Sigma_{i=1}^{N} S_i(t) = C_a(t) + \lambda C_a(t-1)$ under complete markets. This is because $S(t)$ grows at the same rate as as aggregate $S^a(t)$ for each household i and thus $S^i(t+\tau)/S_i(t) = S_a(t+\tau)/S_a(t)$ for each i. Multiplying the numerator and denominator of the aggregate version of equation (10) by

[7] It is easy to show that sufficient conditions for aggregation for cointegration stated in Gonzalo (1993) imply our condition.

[8] See Ogaki (1990) for an aggregation result under complete markets which is applicable to non-homothetic preferences.

$S_a(t)^{-\alpha}$, we see that the asset pricing formula (9) also holds for the aggregate service flow:

$$\frac{E_t[\beta\{S_a(t+1)^{-\alpha}+\beta\lambda S_a(t+2)^{-\alpha}\}R(t+1)]}{E_t[\{S_a(t)^{-\alpha}+\beta\lambda S_a(t+1)^{-\alpha}\}]}=1 \tag{9'}$$

2.3. Insurance Components in Measured Wage Rates

In optimal labour contract models, labour income contains a component that provides workers with some degree of protection against business cycle fluctuations. This insurance component of labour income inserts a wedge between the marginal rate of substitution between leisure and consumption and wages. To utilize information in the first-order condition (6) for estimation and testing, we start from the observation that the cointegration restriction is robust as long as the measured wage rate has the same trend as the marginal rate of substitution. Even when there is a wedge between the real wage rate and the marginal rate of substitution, the stationary restriction holds as long as the insurance component does not have (stochastic or deterministic) trends. Intuition suggests that the fraction of the insurance component in the wage rate is likely to be stationary rather than trending.

In order to formalize this intuition, we consider the following simple model. Let $W^m(t)$ be the measured wage rate and $W(t)$ be the marginal product of labour, which is equated with the marginal rate of substitution as in equation (6). Assume that each firm pays the present value of $W(t+1)[1-l(t+1)]$ in each period t, where the endowment of time is normalized to be one. Using the standard asset pricing formula with aggregate service flows we derived above, the measured labour income of the worker, $y(t)$, is

$$y(t)=\frac{E_t[\beta\{S_a(t+1)^{-\alpha}+\beta\lambda S_a(t+2)^{-\alpha}\}W(t+1)[1-l(t+1)]]}{E_t[S_a(t)^{-\alpha}+\beta\lambda S_a(t+1)^{-\alpha}]} \tag{11}$$

Then the measured real wage rate $W^m(t)$ is $y(t)/[1-l(t)]$. Relation (11) implies that $W^m(t)/W(t)$ is stationary. To see this, divide the left-hand side and right-hand side of relation (11) by $[1-l(t)]W(t)$ and then divide both the numerator and denominator of the right-hand side by $S_a(t)^{-\alpha}$. Then the variables in the right-hand side are stationary. Taking logs, $\ln(W^m(t))-\ln(W(t))$ is stationary. Thus the measured wage rate has the same trend as the marginal product of labour, implying $\ln(W^m(t))-\alpha\ln((C(t))$ is stationary.

3. ESTIMATION AND INFERENCE

In this section we describe a two-step procedure for estimating the consumption curvature parameter and testing the model. Our methods combine Ogaki and Park's (1989) cointegration approach to estimating preference parameters with Hansen and Singleton's (1982) GMM approach.

3.1. Implications of the Cointegration Restriction

The notions of stochastic and deterministic cointegration are useful when the economic variables of interest are modelled as difference stationary with drift.[9] The present paper focuses

[9]The notions of stochatic cointegration and the deterministic cointegration restrictions were defined by Ogaki and Park (1989) and Campbell and Perron (1991). Efficiency gains in estimating the cointegrating vectors from imposing the deterministic cointegration cointegration restriction was discussed by West (1989) for the one-regressor case and by Hansen (1992) and Park (1992) for the general multiple-regressors case.

on processes that are integrated of order one. Suppose that the components of a vector series $X(t)$ are difference stationary with drift. If a linear combination of $X(t)$, $\gamma' X(t)$ is trend stationary, the components of $X(t)$ are said to be (stochastically) cointegrated with a cointegrating vector γ. Consider an additional restriction that the cointegrating vector eliminates the deterministic trends as well as the stochastic trends, so that $\gamma' X(t)$ is stationary. This restriction is called the deterministic cointegration restriction.

We assume that the log of equilibrium consumption is difference stationary with drift.[10] Then the cointegration restriction we derived implies that the log of the real wage rate and the log of non-durable consumption are cointegrated with the deterministic cointegration restriction. The cointegrating vector is $[1, -\alpha]'$.

3.2. Econometric Methodology for Cointegration

This subsection describes the econometric procedure we use. This procedure allows us to test the null hypothesis of stochastic cointegration and the deterministic cointegration restriction.

Let $X(t)$ be a two-dimensional difference stationary process: $X(t) - X(t-1) = \phi + \varepsilon(t)$ for $t \geqslant 1$, where ϕ is a two-dimensional vector of real numbers, where $\varepsilon(t)$ is stationary with mean zero, and where each component of $\varepsilon(t)$ has a positive long-run variance. Suppose that $X(t)$ are cointegrated with a cointegrating vector $(1, -\gamma)$ and that the deterministic cointegration restriction is satisfied. Then we can apply Park's (1992) Canonical Cointegrating Regressions (CCR) procedure[11] to

$$X_1(t) = \theta_c + \gamma X_2(t) + \varepsilon_c(t) \tag{12}$$

This CCR procedure requires us to transform the data before running a regression and corrects for endogeneity and serial correlation. Let $\nu(t) = (\varepsilon_c(t), \varepsilon_2(t))$ where $\varepsilon_2(t)$ is the second element of $\varepsilon(t)$. Define $\Phi(t) = E(\nu(t)\nu(t-i)')$, $\Sigma = \Phi(0)$, $\Gamma = \Sigma_{i=0}^{\infty}\Phi(i)$, and $\Omega = \Sigma_{i=-\infty}^{\infty}\Phi(i)$. Here Ω is the long- run covariance matrix of ν_i. Define

$$\Omega_{11\cdot 2} = \Omega_{11} - \Omega_{12}\Omega_{22}^{-1}\Omega_{21} \tag{13}$$

and $\Gamma_2 = (\Gamma_{12}', \Gamma_{22}')'$, where Ω_{ij} and Γ_{ij} are the ijth component of Ω and Γ, respectively. We make an additional assumption that $\Omega_{11.2}$ is positive. Consider transformations

$$y^*(t) = y(t) + \Pi_y' \nu(t) \tag{14}$$

$$X^*(t) = X(t) + \Pi_x' \nu(t) \tag{15}$$

Because $\nu(t)$ is stationary, $y^*(t)$ and $X^*(t)$ are cointegrated with the same cointegrating vector $(1, -\gamma)$ as $y(t)$ and $X(t)$ for any Π_y and Π_x. The idea of the CCR is to choose Π_y and Π_x so that the OLS estimator is asymptotically efficient when $y^*(t)$ is regressed on $X^*(t)$. This requires

$$\Pi_y = \Sigma^{-1}\Gamma_2\gamma + (0, \Omega_{12}\Omega_{22}^{-1}) \tag{16}$$

$$\Pi_x = \Sigma^{-1}\Gamma_2 \tag{17}$$

In practice, long-run covariance parameters in these formulas are estimated, and estimated Π_y and Π_x are used to transform $y(t)$ and $X(t)$. As long as these parameters are estimated consistently, the resultant CCR estimator is asymptotically efficient.

[10] As shown by Hall (1978), consumption is a random walk when the real interest rate is assumed to be constant. Since we allow the real interest rate to vary over time, the first difference of the log of consumption can have any serial correlation.

[11] See Ogaki (1993a) for a more detailed explanation of CCR-based estimation and testing.

The CCR estimators have asymptotic distributions that can be essentially considered as normal distributions, so that their standard errors can be interpreted in the usual way.[12] An important property of the CCR procedure is that linear restrictions can be tested by χ^2 tests which are free from nuisance parameters. We use χ^2 tests in a regression with spurious deterministic trends added to equation (12) to test for stochastic and deterministic cointegration. For this purpose, the CCR procedure is applied to a regression

$$X_1(t) = \theta_c + \sum_{i=1}^{q} \eta_i t^i + \gamma X_2(t) + \varepsilon_c(t) \tag{18}$$

Let $H(p,q)$ denote the standard Wald statistic to test the hypothesis $\eta_p = \eta_{p+1} = \cdots = \eta_q = 0$ with the estimate of the variance of $\varepsilon_c(t)$ replaced by $\Omega_{11\cdot 2}$ (see Park, 1990, for details). Then $H(p, q)$ converges in distribution to a χ^2_{p-q} random variable under the null of cointegration. In particular, the $H(0, 1)$ statistic tests the deterministic cointegrating restriction. On the other hand, the $H(1, q)$ tests stochastic cointegration.

3.3. The Specification Test Based on the Asset Pricing Equation

Our economic model implies that α, the reciprocal of the long-run intertemporal elasticity of substitution, should eliminate the stochastic trends from consumption and the real wage rate. The model also implies that the same preference parameter enters the asset pricing equation. It is well known that the asset pricing equation based on state- and time-separable utility function fails to explain stock and Treasury bill returns simultaneously (see e.g. Hansen and Singleton, 1982; Mehra and Prescott, 1985). Several authors have found evidence that time-non-separability in preferences could help explain asset returns as we discussed in Section 2.[13]

The econometric model for our GMM procedure is based on the asset pricing equation (9'), which implies $E_t(\varepsilon_g^0(t)) = 0$, where

$$\begin{aligned}\varepsilon_g^0(t) = \beta[(C_a(t+1) + \lambda C_a(t))^{-\alpha} + \lambda\beta(C_a(t+2) + \lambda C_a(t+1))^{-\alpha}]R(t+1) \\ -[(C_a(t) + \lambda C_a(t-1))^{-\alpha} + \lambda\beta(C_a(t+1) + \lambda C_a(t))^{-\alpha}]\end{aligned} \tag{19}$$

where C_a indicates aggregate non-durable consumption. We define $\varepsilon_g(t) = \varepsilon_g^0(t)g(\lambda)/[(1+\beta\lambda)\{C_a(t) + \lambda C_a(t-1)\}^{-\alpha}]$ and $g(\lambda) = 1$ uf $\lambda \leqslant 1$ and $g(\lambda) = 1 + (\lambda - 1)^2$ if $\lambda > 1$. We use $\varepsilon_g(t)$ as the disturbance for the GMM estimation. Since the scale factor $g(\lambda)/[1+\beta\lambda)\{C_a(t) + \lambda C_a(t-1)\}^{-\alpha}]$ is in the information available at t, $E_t(\varepsilon_g(t)) = 0$, We scale the disturbance to achieve stationarity required for the GMM,[14] to avoid the trivial solutions that cause an identification problem, and to incorporate the prior information that λ is likely to be

[12] The CCR estimators are asymptotically efficient, but there are other asymptotic efficient estimators by Phillips and Hansen (1990), Phillips (1991), Saikkonen (1991), and Stock and Watson (1993), among others. Johansen's estimators are often used, but Johansen assumes a Gaussian VAR structure. The CCR does not require this Gaussian VAR assumption, which is important for our purpose because our economic model implies non-linear short-run dynamics. Monte Carlo experiments in Park and Ogaki (1991) show that the CCR estimators have better small sample properties in terms of the mean square error than Johansen's estimators. Following Monte Carlo based recommendations by Park and Ogaki (1991) and Han and Ogaki (1991), we used the prewhitening method and report third-stage CCR estimates and fourth-stage CCR $H(p, q)$ test statistics.

[13] We do not take account of the time aggregation problem that Heaton (1995) studied for an asset pricing model with time-non-separable preferences. We use quarterly data, which cover a longer period than monthly data, to utilize information in the long run needed for cointegrating regressions. Though annual data cover an even longer period, the time aggregation problem for the asset pricing equation can be severe for annual data.

[14] The stationarity assumption of the GMM can be relaxed to some extent, but unit-root non-stationarity is not allowed. Hence the stationary inducing transformation is necessary for our model.

smaller than one in absolute value.[15] Even though the asymptotic theory justifies this type of scaling, small- sample properties of the GMM estimator are affected by the choice of the scaling factor. For this reason, the $g(\lambda)$ function is designed not to affect the disturbance when $\lambda \leq 1$: we have little prior information about which admissible value of λ is more plausible when the absolute value of λ is less than one. The disturbance term is MA of order one because of the time-non-separable specification. The weighting matrix for the GMM estimation must take account of the serial correlation.

A formal test statistic can be formed by using the estimate of α from the cointegrating regression in the GMM procedure to obtain restricted estimates. In this restricted GMM estimation, we estimate only β and λ. We use the same weighting matrix to form unrestricted estimates. We then take the difference of Hansen's (1982) chi-square test (Hansen's J_T test) statistic for the overidentifying restrictions from the restricted estimation and that from the unrestricted estimation, in which β, λ, and α are estimated. The difference is the likelihood ratio type test (denoted by C_T), which has an asymptotic chi-square distribution with one degree of freedom.[16] This two-step procedure does not alter the asymptotic distribution of GMM estimators and test statistics because our cointegrating regression estimator is super-consistent and converges at a rate faster than $T^{1/2}$.

4. DATA

Quarterly seasonally adjusted data in the National Income and Product Accounts (NIPA) were used for consumption. We used three alternative measures of non-durable consumption: non-durables plus services (NDS), non-durables (ND), and food. There are several reasons for looking at food as well as NDS and ND that have been typically used in the literature on the aggregate labour supply model. First, NDS and ND contain durable components such as clothing. Second, using food alone allows us to compare our results with those from the PSID, whose only consumption data are food consumption.

The per capita real consumption series was constructed by dividing the above-mentioned series of constant 1982 dollar consumption by the quarterly average of monthly civilian non-institutional adult (age 16 and over) population. For nominal wages we used a series that is an updated version of that used by Eichenbaum *et al.* That series, average hourly compensation in non-agricultural employment, includes wages, employers' contributions to social security, and other labour income (benefits).[17] Real wages were constructed by dividing nominal wages by the implicit deflator of each of the three consumption measures used. We used the value-weighted average of returns on the New York Stock Exchange and the American Stock Exchange obtained from the Center for Research in Security Prices of the University of Chicago (CRSP) as data for stock returns. Monthly returns for the three months in each quarter were compounded. We used three-month Treasury bill yields in the CRSP risk free file as nominal risk free returns.[18] Real gross returns were obtained from the nominal returns and the implicit

[15] Certain values of λ are not admissible because $C_a(t) + \lambda C_a(t-1)$ cannot be negative. In order to exclude these values in the GMM non-linear search, a very large positive number was returned as $\varepsilon_g(t)$ when they are tried. A numerical derivative program was modified accordingly. See Ogaki (1993b) for details.

[16] See e.g. Ogaki (1993d) for an explanation of the likelihood ratio type test for GMM.

[17] The series used by Eichenbaum *et al.* was a Citibase series, LPCNAG. Since Citibase no longer reports that series, we reconstructed it and updated it using their original definition. The data and the details of its construction are available on request.

[18] Since yields reported in this file are continuously compounded 365 days to maturity, the nominal gross return for our model is $\exp(r \cdot NDM/(365.\ 100))$ where r is the yield reported in percentage and NDM is the number of days to maturity reported in the file.

deflators of each of the three consumption measures. The timing convention for matching returns with consumption are that of Hansen and Singleton (1982). The sample period was from 1947:I to 1990:IV.

5. EMPIRICAL RESULTS

This section reports the results of cointegrating regressions and the specification test based on the asset pricing equations for stock returns and nominal risk free returns.

We first report test results for the null of difference stationarity against trend stationarity. This null usually cannot be rejected for real wage series as in Altonji and Ashenfelter (1980) and Nelson and Plosser (1982) and for consumption series as in Eichenbaum and Hansen (1989) among others. We verify this for our data of real wage and consumption by the $J(1,5)$ test proposed by Park and Choi (1988). Compared with Said and Dickey's (1984) Augmented Dickey–Fuller (ADF) test and Phillips and Perron's (1988) tests, that are often used in the literature, the $J(1,5)$ test has the advantage that neither an estimate of the long-run variance nor a choice of the order of the AR is required. This is very important because the ADF test results are known to be very sensitive to the order of the AR used and because there is no guidance from asymptotic theory to choose the order of the AR. Park and Choi's Monte Carlo experiments show that the $J(1,5)$ test has little size distortion compared with Phillips and Perron tests and is not dominated by the ADF test or Phillips and Perron tests in terms of size adjusted power in small samples.

The values of the $J(1,5)$ test were 5·8, 3·7, 2·3, for the three measures of log real consumption NDS, ND, and food, respectively, and 26·0, 8·9, and 8·7 for the three measures of log real wage rate corresponding to the implicit deflator for NDS, ND, and food, respectively. Thus, we do not reject the null of difference stationarity at 10% level for all series.

Table I. Canonical cointegrating regression results

Regressand (1)	α (2)	$1/\alpha$ (3)	$H(0,1)$ (4)	$H(1,2)$ (5)	$H(1,3)$ (6)	$H(1,4)$ (7)
w, NDS	1·103	0·907	5·644	0·400	7·249	10·805
	(0·075)	(0·094)	(0·018)	(0·527)	(0·027)	(0·013)
c, NDS	1·916	0·522	0·033	1·944	3·039	4·085
	(0·214)	(0·112)	(0·856)	(0·163)	(0·219)	(0·252)
w, ND	2·361	0·424	0·744	2·060	7·243	8·584
	(0·098)	(0·019)	(0·388)	(0·151)	(0·027)	(0·035)
c, ND	2·370	0·422	1·486	0·005	0·032	0·062
	(0·142)	(0·025)	(0·223)	(0·942)	(0·984)	(0·996)
w, Food	2·939	0·340	1·297	1·774	1·808	3·260
	(0·226)	(0·027)	(0·255)	(0·183)	(0·405)	(0·353)
c, Food	2·870	0·348	0·228	0·089	1·446	3·798
	(0·221)	(0·028)	(0·633)	(0·765)	(0·485)	(0·284)

Notes: Park and Ogaki's (1991) method with Andrews' (1991) automatic bandwidth parameter estimator was used to estimate long-run correlation parameters. In column 1, the regressand is indicated as w ($\ln(W)$) or c ($\ln(C)$). The measure of consumption and the corresponding measure of the implicit deflator used to yield real wage rate are indicated as NDS, ND, and food. In columns. 2 and 3, standard errors are in parentheses. Column 4 is a χ^2 test statistic for the deterministic cointegration restriction. Asymptotic P-values are in parentheses. Columns 5, 6 and 7 are χ^2 test statistics for stochastic cointegration. Asymptotic P-values are in parentheses.

Table I reports CCR results.[19] For each measure of consumption, both the log of the real wage rate and the log of real consumption were used as the regressand because either variable can be used in the cointegrating regressions. All point estimates for α have the theoretically correct positive sign. Our estimates indicate that the long-run IES $(1/\alpha)$ is smaller for food, supporting Atkeson and Ogaki's (1991) model.

For ND and food, point estimates are strikingly similar when different regressands are used. This is favourable evidence for cointegration because we expect similar estimates when the

Table II. Specification test results based on the asset pricing equation

Consumption (1)	α (2)	β (3)	λ (4)	J_T (5)	DF (6)	C_T (7)
(1) Using consumption and returns as instruments						
NDS (R)	1·103	0·995	−0·522	9·573	4	...
		(0·003)	(0·120)	(0·048)		
NDS (U)	6·729	1·020	−0·181	6·039	3	3·534
	(2·948)	(0·013)	(0·148)	(0·110)		(0·060)
ND (R)	2·361	0·999	−0·143	7·150	4	...
		(0·002)	(0·089)	(0·128)		
ND (U)	3·092	1·001	−0·101	6·859	3	0·291
	(1·346)	(0·004)	(0·103)	(0·077)		(0·590)
Food (R)	2·939	0·589	−0·958	10·577	4	...
		(0·161)	(0·008)	(0·032)		
Food (U)	2·642	0·607	−0·960	10·545	3	0·032
	(2·988)	(0·360)	(0·017)	(0·014)		(0·858)
(2) Using financial instruments						
NDS (R)	1·103	0·992	−0·773	6·472	2	...
		(0·007)	(0·116)	(0·039)		
NDS (U)	7·129	1·011	−0·475	5·811	1	0·661
	(6·793)	(0·023)	(0·282)	(0·016)		(0·416)
ND (R)	2·361	0·996	−0·284	6·471	2	...
		(0·005)	(0·331)	(0·039)		
ND (U)	4·280	1·000	1·000	6·416	1	0·055
	(3·130)	(0·025)	(267·294)	(0·011)		(0·814)
Food (R)	2·939	0·820	−0·893	3·256	2	...
		(0·100)	(0·047)	(0·196)		
Food (U)	0·447	0·844	−0·971	2·802	1	0·454
	(2·075)	(0·151)	(0·040)	(0·094)		(0·501)

Notes: The estimate of a reported Table I With $\ln(W)$ as the regressand was used for each consuption category for restricted estimation. In column 1 we report the measure of non-durable consumption used and whether the estimation is restricted (R) or unrestriced (U). In columns 2, 3, and 4 standard errors are in parentheses. Column 5 reports Hansen's J-test, and asymptotic P-values are in parentheses. Column 6 is the degree of freedom of Hansen's J in column 5. Column 7 reports the likelihood ratio type test with one degree of freedom for the restriction imposed. The asymptotic p-values are in parentheses.

[19] We used Ogaki's (1993C) GAUSS CCR Package for the CCR estimations. The CCR procedure requires an estimate of the long-run covariance of the disturbances in the system. We used Park and Ogaki's (1991) method with Andrews and Monahan's (1992) prewhitened HAC estimator with the QS kernel. VAR of order one was used for prewhitening. We followed footnote 4 of Andrews and Monahan and the maximum absolute value of the elements of Δ in their notation was set to 0·99. Andrews's (1991) automatic bandwidth estimator, S_T was constructed from fitting AR(1) to each disturbance.

system is cointegrated while there is no reason to expect similarity when the system is not cointegrated. It is not possible, however, to construct a formal test based on this similarity because the estimators with different regressands converge in distribution to the same random variable. The $H(1, q)$ statistics test stochastic cointegration, and we always fail to reject the null of stochastic cointegration at the 1 per cent level. The $H(0, 1)$ tests the deterministic cointegration restriction. We always fail to reject this restriction at the 1 per cent level, and only rejects it at the 5 per cent level for NDS when $\ln(W)$ is used as the regressand.

Table II presents results for the specification test based on the asset pricing equation for the value weighted returns and Treasury bill returns. For each of three measures of non-durable consumption, the estimate of α from the cointegrating regression with the log real wage rate as the regressand was used for the GMM procedure.[20] For the restricted estimation, the parameters estimated were β and λ. For the unrestricted estimation, the parameters estimated were α, b, and λ.

The first panel of Table II reports results when conventional instrumental variables of consumption growth and asset returns are used. We used a constant, real consumption growth lagged one period, and a real gross stock return lagged one period for the disturbance for stock returns; and a constant, real consumption growth lagged one period, and a real gross Treasury bill return lagged one period for the disturbance for Treasury bill returns.

The second panel of Table II reports results when some financial instrumental variables are used. We used a constant, a dividend yield (seasonally adjusted by removing the deterministic seasonal components estimated from a regression with seasonal dummies), a yield spread, and the value weighted return for the disturbance for stock returns and only a constant for the disturbance for Treasury bill returns. The dividend yield is the average dividend yield on the CRSP value-weighted stock index. The yield spread is the monthly yield to maturity of corporate bonds rated Baa by Moody's Investor Services, minus that of the the Aaa corporate yield.

In most cases, we estimate λ to be statistically significantly negative, implying habit formation. The specification test, C_T test, does not reject our model at the 5 per cent level.[21]

6. CONCLUSIONS

In this paper we tested the following three implications of a neoclassical labour contracting model. First, the stochastic trend in the log of real wages is proportional to the stochastic trend in the log of consumption. Second, the cointegrating vector $(1, -\alpha)'$ that eliminates the stochastic trends also eliminates the deterministic trends in the two time series, where α is the reciprocal of the long-run IES for non-durable consumption. Third, the asset pricing equations for stock and nominal risk- free returns are satisfied when the second component of the cointegrating vector is used as the reciprocal of the long-run IES.

---

[20] We used the Hansen–Heaton-Ogaki GAUSS GMM Package that was supported by NSF Grants SES–3512371 and SES–9213930 for the GMM estimation. The initial weighting matrices were identity matrices for each estimation except for estimations with multiple returns, where they were calculated from estimates with stock returns. We iterated on the weighting matrix as described by Kocherlakota (1990) up to five iterations, since his Monte Carlo results indicated that the iteration improves the small sample properties of the GMM estimator. When the zero restrictions lead to a non-positive semidefinite covariance matrix, we used Eichenbaum, Hansen, and Singleton's (1988) modified Durbin's method with AR(5).

[21] We checked sensitivity of results with respect to fixed subsistence levels. All equations used in our econometric work continue to hold with fixed subsistence levels if we replace $C_a(t)$ by $C_a(t) - \gamma$, where γ is Ogaki and Atkeson's (1992) subsistence level estimate of non-durable consumption. We found very little effect on our empirical results from fixed subsistence levels.

We used three alternative measures of non-durable consumption, NDS, ND, and food. We did not reject any of the three implications of our model at the 1 per cent level. Only a few of our specification test statistics were significantly large at the 5 per cent level.

APPENDIX

In this Appendix we argue that the cointegrating coefficient α in the model presented in Section 2 is not constrained to be one even when the ratio of total consumption expenditure and labour income is stationary when a consumer is rich enough. In order to see the relation between this restriction and the curvature parameter α, let us introduce another consumption good, say durable good $D(t)$, explicitly by replacing equation (2) by

$$u(t) = \frac{C(t)^{1-\alpha} - 1}{1-\alpha} + v(l(t)) + \theta \frac{D(t)^{1-\eta} - 1}{1-\eta} \tag{A1}$$

Then the utility function over two consumption goods is Houthakker's (1960) addilog utility function. Let us take the non-durable consumption as a numeraire and let $E(t) = C(t) + P(t)D(t)$ be the total consumption expenditure and $P(t)$ be the relative price.

Atkeson and Ogaki (1991) showed that the intertemporal elasticity of substitution (IES) for the total consumption expenditure is

$$\sigma(t) = \omega(t)\frac{1}{\alpha} + (1 - \omega(t))\frac{1}{\eta} \tag{A2}$$

where $\omega = C/E$ is the budget share for the non-durable consumption. Thus the IES for total consumption expenditure is weighted average of the IES for non-durable consumption, $1/\alpha$, and the IES for durable consumption, $1/\eta$ with the weights being equal to the budget share of each good. For the following discussion, we assume that that $\alpha \geqslant \eta$, then the expenditure elasticity of demand for non-durable consumption is less than or equal to that for durables. This assumption seems reasonable. If $\alpha = \eta$, then preferences are homothetic and the IES is $1/\alpha$. Suppose that $\alpha > \eta$. Then when a consumer is poor, ω is one and the IES is $1/\alpha$.[22] When a consumer becomes rich enough, the IES is approximately constant at $1/\eta$.

Since we assume that leisure is stationary and is additively separable from consumption, the restriction that the ratio of labour income and total consumption expenditure is stationary when a consumer is rich enough implies $\sigma(t)$ is close to one for rich consumers. In order to see this, assume that the relative price $P(t)$ is constant, then

$$W(t) = \frac{v'(l(t))}{AE(t)^{-1/\sigma}} \tag{A3}$$

is satisfied for A that is approximately constant for a rich consumer, and thus $\ln(W(t)) - (1/\sigma)\ln(E(t))$ is stationary. In order for $W(t)/E(t)$ to be stationary for a rich consumer so that the ratio of labour income and total consumption expenditure is stationary, we require that σ is one (see also King, Plosser, and Rebelo, 1988). If preferences are homothetic, then $\alpha = \eta$ and

[22] When a consumer is poor, fixed subsistence levels could be important for the IES as Rebelo (1992, also see Ogaki, 1992b), Atkeson and Ogaki (1991), and Ogaki and Atkeson (1992) discussed. Fixed subsistence levels are not likely to matter for the post-war US aggregate data. Our empirical results turned out to be robust with respect to subsistence levels as we reported above.

$\sigma = \alpha$. Thus this restriction implies $\alpha = 1$. However, if preferences over two consumption goods are not homothetic, α can be different from one. Non-homotheticity implied by $\alpha \neq \eta$ also has important implications for optimal monetary policies (see Braun, 1994).

In our empirical work, we use three alternative measures of non-durable consumption; non-durables plus services (NDS), non-durables (ND), and food. Then durable consumption $D(t)$ in our model should be interpreted as total consumption expenditure minus each measure of non-durable consumption: durables for NDS, durables plus services for ND, and non-food consumption for food. Atkeson and Ogaki (1991) argued that it is more difficult to substitute necessities intertemporally than luxuries, so $1/\alpha$ should be smaller for necessities such as food. Since durables and services are likely to contain more luxuries than ND, we expect $1/\alpha$ to be the smallest for food, and the largest for NDS. Our empirical results reported in Table I are consistent with this view.

ACKNOWLEDGEMENTS

A previous version of this paper was entitled 'Estimating a Labor Supply Model Using Cointegration Restrictions'. We thank Mariarme Baxter, John Campbell, John Cochrane, Robert King, David Marshall, Changyong Rhee, Alan Stockman and seminar participants at Universities of California at San Diego, Houston, Rochester, and Toronto, Arizona State, Rice, and Southern Methodist Universities, and the July 1992 NBER Economic Fluctuations Research Meeting for helpful comments, Jang-Ok Cho for initial collaboration on this project, Hsiang-Ling Han for research assistance, and R. Anton Braun and David Chapman for their help in obtaining data. Special thanks are due to Adrian Pagan, whose comments have improved this paper substantially. This research was supported in part by NSF Grant SES-8921346 and NSF Grant SES-9213930.

REFERENCES

Abowd, J. M., and D. Card (1987), 'Intertemporal labor supply and long-term employment contracts', *American Economic Review*, **77** (March), 50–68.

Allen, E. R. (1992), 'Cross sectional and time series measures of asset pricing model fit', Manuscript, University of Houston.

Altonji, J., and O. Ashenfelter (1980), 'Wage movements and the labour market equilibrium hypothesis', *Economica*, **47** (August), 217–45.

Andrews, D. K. (1991), 'Heteroskedasticity and autocorrelation consistent covariance matrix estimation', *Econometrica* **59** (May), 817–58.

Andrews, D. K., and J. C. Monahan (1992), 'An improved heteroskedasticity and autocorrelation consistent covariance matrix estimator', *Econometrica*, **60** (July), 953–66.

Atkeson, A., and M. Ogaki (1991), 'Wealth-varying intertemporal elasticities of substitution: evidence from panel and aggregate data', Rochester Center for Economic Research Working Paper No. 303, University of Rochester (November).

Azariadis, C. (1975), Implicit contracts and unemployment equilibria', *Journal of Political Economy*, **83** (December), 1183–202.

Benhabib, J., R. Rogerson and R. Wright (1991), 'Homework in macroeconomics: household production and aggregate fluctuations', *Journal of Political Economy*, **99**, 1166–1187.

Braun, R. A. (1994), 'How large is the optimal inflation tax?' *Journal of Monetary Economics*, **34**, 201–214.

Braun, P. A., G. M. Constantinides and W. E. Ferson (1993), 'Time nonseparability in aggregate consumption: international evidence', *European Economic Review*, **37**, 897–920.

Campbell, J. Y., and P. Perron (1991), 'Pitfalls and opportunities: what macroeconomists should know about unit roots,' in O. J. Blanchard and S. Fischer (eds), *NBER Macroeconomics Annual 1991*, Cambridge, MA, 141–210.

Christiano, L. J., and M. Eichenbaum (1990), 'Unit roots in real GNP—do we know and do we care?' *Carnegie-Rochester Conference Series on Public Policy*, **32** (Spring), 7–62.

Cochrane, J. H. (1988), 'How big is the random walk in GNP?' *Journal of Political Economy*, **96** (October), 893–920.

Constantinides, G. M. (1990), 'Habit formation: a resolution of the equity premium puzzle.' *Journal of Political Economy*, **98**, (June), 519–43.

Eichenbaum, M., and L. P. Hansen, (1990), 'Estimating models with intertemporal substitution using aggregate time series data', *Journal of Business & Economic StatisticsTIM*, **8** (January), 53–69.

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

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

Ferson, W. E., and G. M. Constantinides (1991), 'Habit formation and durability in aggregate consumption: empirical tests', *Journal of Financial Economics*, **29**, 199–240.

Gomme, P., and J. Greenwood (1995), 'On the cyclical allocation of risk', *Journal of Economic Dynamics and Control*, **19**, 91–124.

Gonzalo, J. (1993), 'Cointegration and aggregation', *Ricerche-Economiche* **47**, 281–291.

Greenwood, J., and Z. Hercowitz (1991), 'The allocation of goods and time over the business cycle', *Journal of Political Economy*, **99**, 1188–1214.

Hall, R. E. (1978), 'Stochastic implications of the life cycle-permanent income hypothesis: theory and evidence', *Journal of Political Economy*, **86** (December), 971–87.

Hall, R. E. (1980), 'Employment Fluctuations and wage rigidity', *Brooking Papers on Economic Activity*, **1**, 91–123.

Han, H.-L., and M. Ogaki (1991), 'Consumption, income, and cointegration: further analysis', Rochester Center for Economic Research Working Paper No. 305.

Hansen, B. (1992), 'Efficient estimation and testing of cointegrating vectors in the presence of deterministic trends', *Journal of Econometrics*, **53**, 87–121.

Hansen, L. P. (1982), 'Large sample properties of generalized method of moments estimators', *Econometrica*, **50** (July), 1029–54.

Hansen, L· P., and K. J. Singleton (1982), 'Generalized instrumental variable estimation of nonlinear rational expectations models', *Econometrica*, **50** (September), 1269–86.

Hausman, J. A. (1978), 'Specification tests in econometrics', *Econometrica*, **46** (November), 1251–71.

Heaton, J. (1993), 'The interaction between time-nonseparable preferences and time aggregation', *Econometrica*, **61**, 353–85.

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

Hotz, V. J., F. E. Kydland and B. L. Sedlacek (1988), 'Intertemporal preferences and labor supply', *Econometrica*, **56** (March), 335–60.

Houthakker, H. S. (1960), 'Additive preferences', *Econometrica*, **28** (April), 244–56.

Johansen, S. (1988), 'Statistical analysis of cointegration vectors', *Journal of Economic Dynamics and Control*, **12**, 231–54.

Kennan, J. (1988), 'An econometric analysis of fluctuations in aggregate labor supply and demand', *Econometrica*, **56** (March), 317–33.

King, R. G., C. I. Plosser and S. T. Rebelo (1988), 'Production, growth, and business cycles: I. The basic neoclassical model', *Journal of monetary economics*, **21** (March/May), 195–232.

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

Kuznets, S. (1946), *National Income: A Summary of Findings*, National Bureau of Economic Research, New York.

Kydland, F. E. (1980), 'Labor-force heterogeneity and the business cycle', *Carnegie-Rochester Conference Series on Public Policy*, **21**, 173–208.

Kydland, F. E., and E. C. Prescott (1982), 'Time to build and aggregate fluctuations', *Econometrica*, **50** (November), 1345–70.

Mankiw, N. G., J. J. Rotemberg and L. H. Summers (1985), 'Intertemporal substitution in macroeconomics', *Quarterly Journal of Economics*, **C**, 225–52.

Mehra, R., and E. C. Prescott (1985), 'The equity premium: a puzzle', *J. Monetary Econ.*, **15** (January), 145–61.

Nelson, C. R., and C. I. Plosser (1982), 'Trends and random walks in macroeconomic time series', *J. Monetary Econ.*, **10** (September), 139–62.
Ogaki, M. (1990), 'Aggregation of Intratemporal Preferences Under Complete Markets.' Rochester Center for Economic Research Working Paper No. 249, University of Rochester (September).
Ogaki, M. (1992a), 'Engel's law and cointegration', *Journal of Political Economy*, **100**, 1027–46.
Ogaki, M. (1992b), 'Growth in open economies: comment', *Carnegie-Rochester Conference on Public Policy*, **36**, 47–56.
Ogaki, M. (1993a), 'Unit roots in macroeconometrics: a survey', *Monetary and Economic Studies*, **11** (November), 131–54.
Ogaki, M. (1993b), 'GMM: a user's guide', Rochester Center for Economic Research Working Paper No. 348, University of Rochester.
Ogaki, M. (1993c), 'CCR: a user's guide', Rochester Center for Economic Research Working Paper No. 349, University of Rochester.
Ogaki, M. (1993d), 'Generalized method of moments: econometric applications', in G. S. Maddala, C. R. Rao, and H. D. Vinod (eds), *Handbook of Statistics, Vol. 11: Econometrics*, North Holland, Amsterdam, 455–88.
Ogaki, M., and A. Atkeson (1992), 'Estimating subsistence levels with Euler equations from panel data', manuscript, University of Rochester and University of Chicago.
Ogaki, M., and J. Y. Park (1989), 'A cointegration approach to estimating preference parameters', Rochester Center for Economic Research Working Paper No. 209.
Osano, H., and T. Inoue (1991), 'Testing between competing models of real business cycles', *International Economic Review*, **32** (August), 669–88.
Park, J. Y. (1990), 'Testing for unit roots and cointegration by variable addition', *Advances in Econometrics*, **8**, 107–33.
Park, J. Y. (1992), 'Canonical cointegrating regressions', *Econometrica*, **60**, 119–43.
Park, J. Y., and B. Choi (1988), 'A new approach to testing for a unit root', CAE Working Paper No. 88–23, Cornell University.
Park, J. Y., and M. Ogaki (1991), 'Inference in cointegrated models using VAR prewhitening to estimate shortrun dynamics', Rochester Center for Economic Research, Working Paper No. 281, University of Rochester.
Park, J. Y., S. Ouliaris and C. Buhsoo (1988), 'Spurious regressions and tests for cointegration', CAE Working Paper No. 88–07, Cornell University.
Phillips, P. C. B. (1991), 'Optimal inferences in cointegrated systems', *Econometrica*, **59** (March), 283–306.
Phillips, P. C. B., and 'B. E. Hansen (1990), 'Statistical inference in instrumental variables regression with I(1) processes', *Review of Economic Studies*, **57** (January), 99–125.
Phillips, P. C. B., and P. Perron (1988), 'Testing for a unit root in time series regression', *Biometrica*, **75**, 335–46.
Rebelo, S. (1992), 'Growth in open economies', *Carnegie-Rochester Conference Series on Public Policy*, **36**, 5–46.
Rosen, S. (1988), 'Implicit contracts: a survey', *J. Economic Literature*, **23** (September), 1144–75.
Said, S. D., and D. A. Dickey (1984), 'Testing for unit roots in autoregressive-moving average models of unknown order', *Biometrica*, **71**, 599–607.
Saikkonen, P. (1991), 'Asymptotic efficient estimation of cointegration regressions', *Econometric Theory*, **7**, 1–21.
Stock, J. H., and M. W. Watson (1993), 'A simple estimator of cointegrating vectors in higher order integrated systems', *Econometrica*, forthcoming.
Tauchen, G. (1986), 'Statistical properties of generalized-method-of-moments estimators of structural parameters obtained from financial market data', *Journal of Business and Economic Statistics*, **4** (October), 397–425.
West, K. D. (1988), 'Asymptotic normality, when regressors have a unit root', *Econometrica*, **56** (November), 1397–1417.
Wright, R. D. (1988), 'The observational implications of labor contracts in a dynamic general equilibrium model', *Journal of Labor Economics*, **6** (October), 530–51.

LINKED CITATIONS

You have printed the following article:

A Time Series Analysis of Real Wages, Consumption, and Asset Returns
Thomas F. Cooley; Masao Ogaki
Journal of Applied Econometrics, Vol. 11, No. 2. (Mar. - Apr., 1996), pp. 119-134.
Stable URL:
http://links.jstor.org/sici?sici=0883-7252%28199603%2F04%2911%3A2%3C119%3AATSAOR%3E2.0.CO%3B2-4

[Footnotes]

[1] **Intertemporal Labor Supply and Long-Term Employment Contracts**
John M. Abowd; David Card
The American Economic Review, Vol. 77, No. 1. (Mar., 1987), pp. 50-68.
Stable URL:
http://links.jstor.org/sici?sici=0002-8282%28198703%2977%3A1%3C50%3AILSALE%3E2.0.CO%3B2-%23

[2] **Testing between Competing Models of Real Business Cycles**
Hiroshi Osano; Tohru Inoue
International Economic Review, Vol. 32, No. 3. (Aug., 1991), pp. 669-688.
Stable URL:
http://links.jstor.org/sici?sici=0020-6598%28199108%2932%3A3%3C669%3ATBCMOR%3E2.0.CO%3B2-K

[4] **Time to Build and Aggregate Fluctuations**
Finn E. Kydland; Edward C. Prescott
Econometrica, Vol. 50, No. 6. (Nov., 1982), pp. 1345-1370.
Stable URL:
http://links.jstor.org/sici?sici=0012-9682%28198211%2950%3A6%3C1345%3ATTBAAF%3E2.0.CO%3B2-E

[4] **An Econometric Analysis of Fluctuations in Aggregate Labor Supply and Demand**
John Kennan
Econometrica, Vol. 56, No. 2. (Mar., 1988), pp. 317-333.
Stable URL:
http://links.jstor.org/sici?sici=0012-9682%28198803%2956%3A2%3C317%3AAEAOFI%3E2.0.CO%3B2-P

NOTE: *The reference numbering from the original has been maintained in this citation list.*

LINKED CITATIONS

[4] **Intertemporal Preferences and Labor Supply**
V. Joseph Hotz; Finn E. Kydland; Guilherme L. Sedlacek
Econometrica, Vol. 56, No. 2. (Mar., 1988), pp. 335-360.
Stable URL:
http://links.jstor.org/sici?sici=0012-9682%28198803%2956%3A2%3C335%3AIPALS%3E2.0.CO%3B2-W

[4] **A Time Series Analysis of Representative Agent Models of Consumption and Leisure Choice under Uncertainty**
Martin S. Eichenbaum; Lars Peter Hansen; Kenneth J. Singleton
The Quarterly Journal of Economics, Vol. 103, No. 1. (Feb., 1988), pp. 51-78.
Stable URL:
http://links.jstor.org/sici?sici=0033-5533%28198802%29103%3A1%3C51%3AATSAOR%3E2.0.CO%3B2-V

[5] **Habit Formation: A Resolution of the Equity Premium Puzzle**
George M. Constantinides
The Journal of Political Economy, Vol. 98, No. 3. (Jun., 1990), pp. 519-543.
Stable URL:
http://links.jstor.org/sici?sici=0022-3808%28199006%2998%3A3%3C519%3AHFAROT%3E2.0.CO%3B2-T

[9] **Canonical Cointegrating Regressions**
Joon Y. Park
Econometrica, Vol. 60, No. 1. (Jan., 1992), pp. 119-143.
Stable URL:
http://links.jstor.org/sici?sici=0012-9682%28199201%2960%3A1%3C119%3ACCR%3E2.0.CO%3B2-R

[10] **Stochastic Implications of the Life Cycle-Permanent Income Hypothesis: Theory and Evidence**
Robert E. Hall
The Journal of Political Economy, Vol. 86, No. 6. (Dec., 1978), pp. 971-987.
Stable URL:
http://links.jstor.org/sici?sici=0022-3808%28197812%2986%3A6%3C971%3ASIOTLC%3E2.0.CO%3B2-K

[12] **Statistical Inference in Instrumental Variables Regression with I(1) Processes**
Peter C. B. Phillips; Bruce E. Hansen
The Review of Economic Studies, Vol. 57, No. 1. (Jan., 1990), pp. 99-125.
Stable URL:
http://links.jstor.org/sici?sici=0034-6527%28199001%2957%3A1%3C99%3ASIIIVR%3E2.0.CO%3B2-6

NOTE: *The reference numbering from the original has been maintained in this citation list.*

LINKED CITATIONS

[12] **Optimal Inference in Cointegrated Systems**
P. C. B. Phillips
Econometrica, Vol. 59, No. 2. (Mar., 1991), pp. 283-306.
Stable URL:
http://links.jstor.org/sici?sici=0012-9682%28199103%2959%3A2%3C283%3AOIICS%3E2.0.CO%3B2-5

[12] **A Simple Estimator of Cointegrating Vectors in Higher Order Integrated Systems**
James H. Stock; Mark W. Watson
Econometrica, Vol. 61, No. 4. (Jul., 1993), pp. 783-820.
Stable URL:
http://links.jstor.org/sici?sici=0012-9682%28199307%2961%3A4%3C783%3AASEOCV%3E2.0.CO%3B2-Z

[13] **An Empirical Investigation of Asset Pricing with Temporally Dependent Preference Specifications**
John Heaton
Econometrica, Vol. 63, No. 3. (May, 1995), pp. 681-717.
Stable URL:
http://links.jstor.org/sici?sici=0012-9682%28199505%2963%3A3%3C681%3AAEIOAP%3E2.0.CO%3B2-L

[19] **An Improved Heteroskedasticity and Autocorrelation Consistent Covariance Matrix Estimator**
Donald W. K. Andrews; J. Christopher Monahan
Econometrica, Vol. 60, No. 4. (Jul., 1992), pp. 953-966.
Stable URL:
http://links.jstor.org/sici?sici=0012-9682%28199207%2960%3A4%3C953%3AAIHAAC%3E2.0.CO%3B2-Q

[19] **Heteroskedasticity and Autocorrelation Consistent Covariance Matrix Estimation**
Donald W. K. Andrews
Econometrica, Vol. 59, No. 3. (May, 1991), pp. 817-858.
Stable URL:
http://links.jstor.org/sici?sici=0012-9682%28199105%2959%3A3%3C817%3AHAACCM%3E2.0.CO%3B2-P

[20] **A Time Series Analysis of Representative Agent Models of Consumption and Leisure Choice under Uncertainty**
Martin S. Eichenbaum; Lars Peter Hansen; Kenneth J. Singleton
The Quarterly Journal of Economics, Vol. 103, No. 1. (Feb., 1988), pp. 51-78.
Stable URL:
http://links.jstor.org/sici?sici=0033-5533%28198802%29103%3A1%3C51%3AATSAOR%3E2.0.CO%3B2-V

NOTE: *The reference numbering from the original has been maintained in this citation list.*

LINKED CITATIONS

References

Intertemporal Labor Supply and Long-Term Employment Contracts
John M. Abowd; David Card
The American Economic Review, Vol. 77, No. 1. (Mar., 1987), pp. 50-68.
Stable URL:
http://links.jstor.org/sici?sici=0002-8282%28198703%2977%3A1%3C50%3AILSALE%3E2.0.CO%3B2-%23

Wage Movements and the Labour Market Equilibrium Hypothesis
Joseph Altonji; Orley Ashenfelter
Economica, New Series, Vol. 47, No. 187, Special Issue on Unemployment. (Aug., 1980), pp. 217-245.
Stable URL:
http://links.jstor.org/sici?sici=0013-0427%28198008%292%3A47%3A187%3C217%3AWMATLM%3E2.0.CO%3B2-%23

Heteroskedasticity and Autocorrelation Consistent Covariance Matrix Estimation
Donald W. K. Andrews
Econometrica, Vol. 59, No. 3. (May, 1991), pp. 817-858.
Stable URL:
http://links.jstor.org/sici?sici=0012-9682%28199105%2959%3A3%3C817%3AHAACCM%3E2.0.CO%3B2-P

An Improved Heteroskedasticity and Autocorrelation Consistent Covariance Matrix Estimator
Donald W. K. Andrews; J. Christopher Monahan
Econometrica, Vol. 60, No. 4. (Jul., 1992), pp. 953-966.
Stable URL:
http://links.jstor.org/sici?sici=0012-9682%28199207%2960%3A4%3C953%3AAIHAAC%3E2.0.CO%3B2-Q

Implicit Contracts and Underemployment Equilibria
Costas Azariadis
The Journal of Political Economy, Vol. 83, No. 6. (Dec., 1975), pp. 1183-1202.
Stable URL:
http://links.jstor.org/sici?sici=0022-3808%28197512%2983%3A6%3C1183%3AICAUE%3E2.0.CO%3B2-B

LINKED CITATIONS

How Big Is the Random Walk in GNP?
John H. Cochrane
The Journal of Political Economy, Vol. 96, No. 5. (Oct., 1988), pp. 893-920.
Stable URL:
http://links.jstor.org/sici?sici=0022-3808%28198810%2996%3A5%3C893%3AHBITRW%3E2.0.CO%3B2-P

Habit Formation: A Resolution of the Equity Premium Puzzle
George M. Constantinides
The Journal of Political Economy, Vol. 98, No. 3. (Jun., 1990), pp. 519-543.
Stable URL:
http://links.jstor.org/sici?sici=0022-3808%28199006%2998%3A3%3C519%3AHFAROT%3E2.0.CO%3B2-T

A Time Series Analysis of Representative Agent Models of Consumption and Leisure Choice under Uncertainty
Martin S. Eichenbaum; Lars Peter Hansen; Kenneth J. Singleton
The Quarterly Journal of Economics, Vol. 103, No. 1. (Feb., 1988), pp. 51-78.
Stable URL:
http://links.jstor.org/sici?sici=0033-5533%28198802%29103%3A1%3C51%3AATSAOR%3E2.0.CO%3B2-V

Co-Integration and Error Correction: Representation, Estimation, and Testing
Robert F. Engle; C. W. J. Granger
Econometrica, Vol. 55, No. 2. (Mar., 1987), pp. 251-276.
Stable URL:
http://links.jstor.org/sici?sici=0012-9682%28198703%2955%3A2%3C251%3ACAECRE%3E2.0.CO%3B2-T

The Allocation of Capital and Time over the Business Cycle
Jeremy Greenwood; Zvi Hercowitz
The Journal of Political Economy, Vol. 99, No. 6. (Dec., 1991), pp. 1188-1214.
Stable URL:
http://links.jstor.org/sici?sici=0022-3808%28199112%2999%3A6%3C1188%3ATAOCAT%3E2.0.CO%3B2-O

Stochastic Implications of the Life Cycle-Permanent Income Hypothesis: Theory and Evidence
Robert E. Hall
The Journal of Political Economy, Vol. 86, No. 6. (Dec., 1978), pp. 971-987.
Stable URL:
http://links.jstor.org/sici?sici=0022-3808%28197812%2986%3A6%3C971%3ASIOTLC%3E2.0.CO%3B2-K

NOTE: *The reference numbering from the original has been maintained in this citation list.*

LINKED CITATIONS

Large Sample Properties of Generalized Method of Moments Estimators
Lars Peter Hansen
Econometrica, Vol. 50, No. 4. (Jul., 1982), pp. 1029-1054.
Stable URL:
http://links.jstor.org/sici?sici=0012-9682%28198207%2950%3A4%3C1029%3ALSPOGM%3E2.0.CO%3B2-O

Generalized Instrumental Variables Estimation of Nonlinear Rational Expectations Models
Lars Peter Hansen; Kenneth J. Singleton
Econometrica, Vol. 50, No. 5. (Sep., 1982), pp. 1269-1286.
Stable URL:
http://links.jstor.org/sici?sici=0012-9682%28198209%2950%3A5%3C1269%3AGIVEON%3E2.0.CO%3B2-G

Specification Tests in Econometrics
J. A. Hausman
Econometrica, Vol. 46, No. 6. (Nov., 1978), pp. 1251-1271.
Stable URL:
http://links.jstor.org/sici?sici=0012-9682%28197811%2946%3A6%3C1251%3ASTIE%3E2.0.CO%3B2-X

The Interaction Between Time-Nonseparable Preferences and Time Aggregation
John Heaton
Econometrica, Vol. 61, No. 2. (Mar., 1993), pp. 353-385.
Stable URL:
http://links.jstor.org/sici?sici=0012-9682%28199303%2961%3A2%3C353%3ATIBTPA%3E2.0.CO%3B2-0

An Empirical Investigation of Asset Pricing with Temporally Dependent Preference Specifications
John Heaton
Econometrica, Vol. 63, No. 3. (May, 1995), pp. 681-717.
Stable URL:
http://links.jstor.org/sici?sici=0012-9682%28199505%2963%3A3%3C681%3AAEIOAP%3E2.0.CO%3B2-L

Intertemporal Preferences and Labor Supply
V. Joseph Hotz; Finn E. Kydland; Guilherme L. Sedlacek
Econometrica, Vol. 56, No. 2. (Mar., 1988), pp. 335-360.
Stable URL:
http://links.jstor.org/sici?sici=0012-9682%28198803%2956%3A2%3C335%3AIPALS%3E2.0.CO%3B2-W

Additive Preferences
H. S. Houthakker
Econometrica, Vol. 28, No. 2. (Apr., 1960), pp. 244-257.
Stable URL:
http://links.jstor.org/sici?sici=0012-9682%28196004%2928%3A2%3C244%3AAP%3E2.0.CO%3B2-K

An Econometric Analysis of Fluctuations in Aggregate Labor Supply and Demand
John Kennan
Econometrica, Vol. 56, No. 2. (Mar., 1988), pp. 317-333.
Stable URL:
http://links.jstor.org/sici?sici=0012-9682%28198803%2956%3A2%3C317%3AAEAOFI%3E2.0.CO%3B2-P

Time to Build and Aggregate Fluctuations
Finn E. Kydland; Edward C. Prescott
Econometrica, Vol. 50, No. 6. (Nov., 1982), pp. 1345-1370.
Stable URL:
http://links.jstor.org/sici?sici=0012-9682%28198211%2950%3A6%3C1345%3ATTBAAF%3E2.0.CO%3B2-E

Intertemporal Substitution in Macroeconomics
N. Gregory Mankiw; Julio J. Rotemberg; Lawrence H. Summers
The Quarterly Journal of Economics, Vol. 100, No. 1. (Feb., 1985), pp. 225-251.
Stable URL:
http://links.jstor.org/sici?sici=0033-5533%28198502%29100%3A1%3C225%3AISIM%3E2.0.CO%3B2-2

Engel's Law and Cointegration
Masao Ogaki
The Journal of Political Economy, Vol. 100, No. 5. (Oct., 1992), pp. 1027-1046.
Stable URL:
http://links.jstor.org/sici?sici=0022-3808%28199210%29100%3A5%3C1027%3AELAC%3E2.0.CO%3B2-D

Testing between Competing Models of Real Business Cycles
Hiroshi Osano; Tohru Inoue
International Economic Review, Vol. 32, No. 3. (Aug., 1991), pp. 669-688.
Stable URL:
http://links.jstor.org/sici?sici=0020-6598%28199108%2932%3A3%3C669%3ATBCMOR%3E2.0.CO%3B2-K

NOTE: *The reference numbering from the original has been maintained in this citation list.*

Canonical Cointegrating Regressions
Joon Y. Park
Econometrica, Vol. 60, No. 1. (Jan., 1992), pp. 119-143.
Stable URL:
http://links.jstor.org/sici?sici=0012-9682%28199201%2960%3A1%3C119%3ACCR%3E2.0.CO%3B2-R

Optimal Inference in Cointegrated Systems
P. C. B. Phillips
Econometrica, Vol. 59, No. 2. (Mar., 1991), pp. 283-306.
Stable URL:
http://links.jstor.org/sici?sici=0012-9682%28199103%2959%3A2%3C283%3AOIICS%3E2.0.CO%3B2-5

Statistical Inference in Instrumental Variables Regression with I(1) Processes
Peter C. B. Phillips; Bruce E. Hansen
The Review of Economic Studies, Vol. 57, No. 1. (Jan., 1990), pp. 99-125.
Stable URL:
http://links.jstor.org/sici?sici=0034-6527%28199001%2957%3A1%3C99%3ASIIIVR%3E2.0.CO%3B2-6

A Simple Estimator of Cointegrating Vectors in Higher Order Integrated Systems
James H. Stock; Mark W. Watson
Econometrica, Vol. 61, No. 4. (Jul., 1993), pp. 783-820.
Stable URL:
http://links.jstor.org/sici?sici=0012-9682%28199307%2961%3A4%3C783%3AASEOCV%3E2.0.CO%3B2-Z

Statistical Properties of Generalized Method-of-Moments Estimators of Structural Parameters Obtained from Financial Market Data
George Tauchen
Journal of Business & Economic Statistics, Vol. 4, No. 4. (Oct., 1986), pp. 397-416.
Stable URL:
http://links.jstor.org/sici?sici=0735-0015%28198610%294%3A4%3C397%3ASPOGME%3E2.0.CO%3B2-Z

Asymptotic Normality, When Regressors Have a Unit Root
Kenneth D. West
Econometrica, Vol. 56, No. 6. (Nov., 1988), pp. 1397-1417.
Stable URL:
http://links.jstor.org/sici?sici=0012-9682%28198811%2956%3A6%3C1397%3AANWRHA%3E2.0.CO%3B2-K

The Observational Implications of Labor Contracts in a Dynamic General Equilibrium Model
Randall D. Wright
Journal of Labor Economics, Vol. 6, No. 4. (Oct., 1988), pp. 530-551.
Stable URL:
http://links.jstor.org/sici?sici=0734-306X%28198810%296%3A4%3C530%3ATOIOLC%3E2.0.CO%3B2-S