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Working Paper No. 285
June 1991

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First Draft: December 1990

This paper will be presented at the NBER Economic Fluctuations Research Meeting on July 12, 1991 in Boston. A previous version of this paper was entitled "Estimating a Labor Supply Model Using Cointegration Restrictions." We thank Marianne Baxter, John Cochrane, Robert King, David Marshall, Changyong Rhee, Alan Stockman and seminar participants at University of Rochester 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.

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Abstract

This paper reexamines whether or not 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 labor 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 for estimating and testing the model, using Ogaki and Park's (1990) cointegration approach. This long-run restriction involves a parameter that we call the long-run intertemporal elasticity of substitution (IES) for nondurable consumption but does not involve the IES for leisure. This allows us to estimate the long-run IES for nondurable consumption from a cointegrating regression. In the second step of our analysis, our estimates of the long-run IES for nondurable consumption are used to estimate the discount factor and a coefficient of time-nonseparability using Hansen's (1982) Generalized Method of Moments. We allow for time-nonseparability, so that the model could explain both stock and nominal risk fee returns. We form a specification test for our model *a la* Hausman (1978) from these two steps. We do not reject the model and obtain reasonable estimates of preference parameters.

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I. Introduction

This paper reexamines 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 labor contracting model is more appropriate for understanding the time series behavior of real wages and consumption. We show that an optimal contract model restricts the long-run relation between real wages and consumption. We exploit this long-run restriction to estimate preference parameters and test the model using a two step procedure. In the first step we employ the cointegration approach suggested by Ogaki and Park (1990) to estimate the long-run intertemporal elasticity of substitution for nondurable consumption from a cointegrating regression. We use this estimated preference parameter in the asset pricing equation implied by this economy and then estimate the discount factor and a coefficient of time-nonseparability using Hansen's (1982) Generalized Method of Moments (GMM). From this we are able to construct a specification test of the model. The results of this procedure are extremely encouraging. We obtain reasonable estimates of the preference parameters and we do not reject the model.[1]

Mankiw, Rotemberg, and Summers (1986, hereafter Mankiw et al.) subjected the Euler equations of a version of an intertemporal labor supply

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

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.[2] Second, Eichenbaum et al. allowed time-nonseparability of preferences. Though their point estimates were reasonable, the formal test statistics typically rejected the model at the one percent 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 labor contract model might be appropriate for understanding real wages.[3]

A given Pareto optimal allocation can be consistent with a wide variety

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[2]The asymptotic theory for the GMM requires the variables used to be stationary.

[3]Osano and Inoue (1988) 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 rate of substitution.

of institutional arrangements. In optimal labor contract models (see, e.g, Azariades (1978), Rosen (1985), and Wright (1988)), labor income contains a component that provides workers with some degree of protection against business cycle fluctuations. This insurance component of labor income inserts a wedge between the marginal rate of substitution between leisure and consumption and wages. In their empirical work, Gomme and Greenwood (1990) 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 labor 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 the optimal contract model does restrict the long-run relation of the real wage rate and consumption. 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 nondurable consumption. We do not reject the null cf 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 the GMM to estimate the discount parameter and a coefficient of time-nonseparability. We use both stock and nominal risk free returns. We form a specification test *a la* Hausman (1978) through these steps. This specification test does not reject the model and we obtain reasonable estimates of the discount factor and a coefficient of time-nonseparability.

The rest of the present paper is organized as follows. In Section II, we present our model and derive the cointegration restriction. We describe our econometric procedures in Section III. In Section IV, we explain the data. Section V contains our empirical results. Our concluding remarks are in Section VI.

II. The Economy

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. Household i maximizes

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

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

$$u_i(t) = \frac{C_i^n(t)^{1-\alpha} - 1}{1-\alpha} + v_i^1(\ell_i(t)) + v_i^d(C_i^d(t)) \tag{2}$$

where $v_i^1(\cdot)$ is a continuously differentiable concave function, $C_i^n(t)$ is nondurable consumption, $C_i^d(t)$ is durable consumption, and $\ell_i(t)$ is leisure of consumer i.

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

$$W_i(t) = \frac{v_i'(\ell_i(t))}{C_i(t)^{-\alpha}}, \tag{3}$$

where $W_i(t)$ is the real wage rate for worker i. We assume that the stochastic process of leisure is (strictly) stationary in the equilibrium as in Eichenbaum, Hansen, and Singleton 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 $\ln(W_i(t)) - \alpha\ln(C_i(t)) = \ln(v_i'(\ell_i(t))$ 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 U.S. as Kuznets (1946) found, it is natural to impose a restriction that the ratio of total consumption expenditure and labor income is stable when a consumer is rich enough. In order to see the relation between this restriction and the curvature parameter α, let us assume that

$$v_i^d(C_i^d(t)) = \theta_2 \frac{C_i^d(t)^{1-\eta} - 1}{1-\eta} \tag{4}.$$

Then the utility function over two consumption goods is Houthakker's (1960) addilog utility function. Let us take the nondurable consumption as a numeraire and let $E_i(t)=C_i^n(t) + P(t)C_i^d(t)$ 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

$$\sigma_i(t) = \omega_i^n(t)\frac{1}{\alpha} + (1-\omega_i^n(t)) \frac{1}{\eta} \tag{5}$$

where $\omega_i^n = C_i^n/E_i$ is the budget share for the nondurable consumption. Thus the IES for total consumption expenditure is weighted average of the IES for nondurable 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 $\alpha\geq\eta$, then the expenditure elasticity of demand for nondurable 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_i^n is one and the IES is $1/\alpha$.[4] 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 labor income and total consumption expenditure is stationary when a consumer is rich enough implies $\sigma_i(t)$ is close to one for rich consumers. In order to see this, assume that the relative price P is constant, then

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[4]When a consumer is poor, fixed subsistence levels could be important for the IES as Blundell, Browning, and Meghir (1991), Rebelo (1991, also see Ogaki (1991)), Atkeson and Ogaki (1991), and Ogaki and Atkeson (1991b) discussed. Fixed subsistence levels are not likely to matter for the post war U.S. aggregate data. Our empirical results turn out to be robust with respect to subsistence levels as we report below.

$$W_i(t) = \frac{v_i'(\ell_i(t))}{AE_i(t)^{-1/\sigma}}, \tag{6}$$

is satisfied for A that is approximately constant for a rich consumer, and thus $\ln(W)-(1/\sigma)\ln(E)$ is stationary. In order for W/E to be stationary for a rich consumer so that the ratio of labor 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 $\sigma=\alpha$. Thus this restriction implies $\alpha=1$. However, if preferences over two consumption goods are not homothetic, α can be different from one. Nonhomotheticity implied by $\alpha\neq\eta$ also has important implications for optimal monetary policies (see Braun (1991)).

In our empirical work, we use three alternative measures of nondurable consumption; nondurables plus services (NDS), nondurables (ND), and food. Then durable consumption C_i^d in our model should be interpreted as total consumption expenditure minus each measure of nondurable consumption: durables for NDS, durables plus services for ND, and nonfood 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.

We now introduce time-nonseparability of preferences. The intraperiod utility function is assumed to be

$$u_i(t) = \frac{S_i^n(t)^{1-\alpha} - 1}{1-\alpha} + v_i^1(S_i^1(t)) + v_i^d(C_i^d(t),\ldots,\ C_i^d(t-k)) \tag{7}$$

where $v_i^1(\cdot)$ and $v_i^n(\cdot)$ are continuously differentiable concave functions and

$S^n_i(t)$, $S^l_i(t)$, and are the service flows from nondurable consumption and leisure, respectively:

$$S^n_i(t) = C_i(t) + \lambda^n C_i(t-1) \qquad (8),$$

and

$$S^l_i(t) = \ell_i(t) + \lambda^l_1 \ell_i(t-1) + \ . \ . \ . + \lambda^l_k \ell_i(t-k) \qquad (9).$$

This service flow specification of leisure has been used by many authors[5] 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 (1990)).

The time-nonseparable specification for nondurable consumption is similar to that considered by Eichenbaum, Hansen, and Singleton (1988), Eichenbaum and Hansen (1990), Constantinides (1990), and Heaton (1990, 1991) among others, except that some of these authors considered more general form of time-nonseparability for nondurable consumption than (8). We assume $\lambda^n > -1$, so that marginal utility of consumption with respect to the life-time utility is positive. We have habit formation for nondurable consumption when λ^n is negative and local substitutability or durablity when λ^n is negative. The time-nonseparability for nondurable 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

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

(1991) found evidence in favor of the asset pricing model with habit formation, using the GMM procedure. Note that the time-nonseparability does not affect the IES in the long-run when $C_i^n(t)$ and $C_i^n(t-1)$ are equal.[6] We will refer to $1/\alpha$ as the long-run IES for nondurable 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:

$$W_i(t) = \frac{\partial U_i/\partial \ell_i(t)}{\partial U_i/\partial C_i^n(t)} = \frac{E_t[\sum_{\tau=0}^{K} \beta^\tau \partial u(t+\tau)/\partial \mathcal{L}_i(t)]}{E_t[\partial u_i(t)/\partial C_i^n(t) + \partial u_i(t+1)/\partial C_i^n(t)]}$$

$$= \frac{E_t[\sum_{\tau=0}^{K} \beta^\tau \lambda_\tau^i v_i'(S_i^1(t+\tau))]}{E_t[S_i^n(t)^{-\alpha} + \beta\lambda^n S_i^n(t+1)^{-\alpha}]} \quad (10).$$

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

$$S_i(t+\tau)/C_i(t) = C_i(t+\tau)/C_i(t) + \lambda^n C_i(t+\tau-1)/C_i(t) \quad (11)$$

is stationary for any τ. Combined with the first order condition (11), this implies that

$$W_i(t)C_i(t)^{-\alpha} = \frac{E_t[\sum_{\tau=0}^{K} \beta^\tau \sigma_i b_\tau^i v_i'(S_i^1(t+\tau))]}{E_t[\{S_i^n(t)/C_i^n(t)\}^{-\alpha} + \{S_i^n(t+1)/C_i^n(t)\}^{-\alpha}]} \quad (12)$$

is stationary. Taking logs, $\ln(W_i(t)) - \alpha\ln(C_i(t))$ is stationary as in the time-separable case we discussed.

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[6] Alternatively, $C_i^n(t)$ grows at a constant rate in the long-run.

Aggregation

We have derived the cointegration restriction for individual households. This restriction also holds for aggregated data under certain conditions. Let $W_a = (1/N)\sum_{i=1}^{N} W_i$ and $C^n_a = (1/N)\sum_{i=1}^{N} C^n_i$. The cointegration restriction implies $(1/N)\sum_{i=1}^{N}\log(W_i) - \alpha(1/N)\sum_{i=1}^{N}\log(C^n_i)$ 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 service grows at the same rate for all consumers under the perfect risk sharing and because the ratio of service flow and 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 (10) through the cointegration restriction for aggregated real wages and consumption. We also estimate and test the standard asset pricing equation for the time-nonseparable utility function

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[7]It is easy to see that sufficient conditions for aggregation for cointegration stated in Appendix of Gonzalo (1989) imply our condition.

[8]See Ogaki (1990) for an aggregation result under complete markets which is applicable to nonhomothetic preferences and Ogaki and Atkeson (1991b) and references cited therein for implications of risk sharing.

$$\frac{E_t[\beta\{(S_i^n(t+1)-\gamma_s)^{-\alpha}+\beta\lambda^n(S_i^n(t+2))\}^{-\alpha}R(t+1)]}{E_t[\{(S_i^n(t))^{-\alpha}+\beta\lambda^n(S_i^n(t+1))\}^{-\alpha}]} = 1 \tag{13}$$

for any gross asset return $R(t)$. Dividing both the numerator and denominator of (13) by $S_i^n(t)^{-\alpha}$, which is in the information set available at t, we obtain

$$\frac{E_t[\beta\{(S_i^n(t+1)/S_i^n(t))^{-\alpha}+\beta\lambda^n(S_i^n(t+2)/S_i^n(t))^{-\alpha}\}R(t+1)]}{E_t[1+\beta\lambda^n(S_i^n(t+1)/S_i^n(t))^{-\alpha}]} =1 \tag{14}$$

This asset pricing formula is also satisfied by the aggregated service flow $S_a^n(t) \equiv (1/N)\sum_{i=1}^{N} S_i^n = C_a^n(t)+\lambda^n C_a(t-1)$ under complete markets. This is because $S_i^n(t)$ grows at the same rate as as aggregate $S_a^n(t)$ for each i and thus $S_i^n(t+\tau)/S_i^n(t)=S_a^n(t+\tau)/S_a^n(t)$ for each i. Multiplying the numerator and denominator of the aggregate version of (14) by $S_a^n(t)^{-\alpha}$, we see that the asset pricing formula (13) also holds for the aggregate service flow.

Measured Wage Rates

In optimal labor contract models, labor income contains a component that provides workers with some degree of protection against business cycle fluctuations. This insurance component of labor income inserts a wedge between the marginal rate of substitution between leisure and consumption and wages. In order to utilize information in the first order condition (5) 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. Intuitively, the insurance component is likely to be stationary in nature rather than trending.

In order to formalize this intuition, we consider the following simple model. Let $W_i^m(t)$ be the measured wage rate and $W_i(t)$ be the marginal product of labor, which is equated with the marginal rate of substitution as in (5). Assume that each firm pays the present value of $W_i(t+1)[1-\ell_i(t+1)]$ 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, the measured labor income of worker i, $y_i^1(t)$, is

$$y_i^1(t)=\frac{E_t[\beta\{S_a^n(t+1)^{-\alpha}+\beta\lambda^n S_a^n(t+2)^{-\alpha}\}W_i(t+1)[1-\ell_i(t+1)]}{E_t[S_a^n(t)^{-\alpha}+\beta\lambda^n S_a^n(t+1)^{-\alpha}]} \tag{15}$$

Then the measured real wage rate $W_i^m(t)$ is $y_i(t)/[1-\ell_i(t)]$. Relation (15) implies that $W_i^m(t)/W_i(t)$ is stationary. To see this, divide the left hand side and right hand side of (15) by $[1-\ell_i(t)]W_i(t)$ 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, $\ln(W_i^m(t))-\ln(W_i(t))$ is stationary. Thus the measured wage rate has the same trend as the marginal product of labor, implying $\ln(W_i^m(t))-\alpha\ln(C_i(t))$ is stationary.

III. 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 (1990) cointegration approach to estimating preference parameters with Hansen and Singleton's (1982) GMM procedure.

A. Implications of the Cointegration Restriction

Ogaki and Park's (1990) notions of stochastic and deterministic cointegration are useful when the economic variables of interest are modeled as difference stationary with drift (also see Campbell and Perron (1991)). The present paper focuses 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.[9] Then the cointegration restriction we derived implies that the log of the real wage rate and the log of nondurable consumption are cointegrated with the deterministic cointegration restriction. The cointegrating vector is $[1, -\alpha]'$.

C. Econometric Methodology for Cointegration

This subsection briefly describes the econometric procedure used by Ogaki and Park (1990) for cointegrated systems. This procedure allows us to test the null of deterministic cointegration restriction and stochastic cointegration.

Let $X(t)$ be a 2-dimensional difference stationary process: $X(t) - X(t-1) = \phi + \epsilon(t)$ for $t \geq 1$, where ϕ is a 2-dimensional vector of real numbers, where $\epsilon(t)$ is stationary with mean zero, and where each component

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[9] 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.

of $\epsilon(t)$ 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 the Canonical Cointegrating Regressions (*CCR*) procedure developed by Park (1990) to

$$X_1(t) = \theta_c + \gamma X_2(t) + \epsilon_c(t), \tag{16}$$

This *CCR* procedure requires us to transform the data before running a regression and corrects for endogeneity and serial correlation. 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.[10]

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 (16) to test for stochastic and deterministic cointegration. For this purpose, the *CCR* procedure is applied to a regression

$$X_1(t) = \theta_c + \sum_{i=1}^{q} \eta_i t^i + \gamma X_2(t) + \epsilon_c(t). \tag{17}$$

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[10] The CCR estimators are estimators are asymptotically efficient, but there are other asymptotic efficient estimators by Phillips (1988a, 1988b), Stock and Watson (1989), Saikkonen (1989), and Phillips and Hansen (1990) among others. Johansen's estimators are often used, but Johansen assumes Gaussian VAR structure. The CCR does not require this Gaussian VAR assumption, which is important for our purpose because our economic model implies nonlinear 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. For small sample efficiency, we used Park and Ogaki's VAR prewhitening method and the second step CCR estimator, which is formed with long-run correlation parameters estimated by the first step CCR based on OLS as their Monte Carlo experiments suggested.

Let $H(p,q)$ denote the standard Wold statistic to test the hypothesis $\eta_p=\eta_{p+1}= \ldots =\eta_q=0$ with the estimate of the variance of $\epsilon_c(t)$ replaced by the long run variance of the *CCR* (see Park (1990) for more explanation). 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.

The CCR allows to test the null hypothesis of the deterministic cointegration restriction as well as stochastic cointegration.[11] It is important to test the deterministic cointegration restriction because it is difficult to discriminate between a unit root process and a stationary process in finite samples (see, e.g., Cochrane (1988), Christiano and Eichenbaum (1989), and Campbell and Perron (1991)). This means that stochastic cointegration is hard to test.[12]

D. 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

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[11]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 (1990) and Park (1990) for the general multiple regressors case.

[12]Preliminary Monte Carlo results in Han and Ogaki (1991) suggest that the H(0,1) test for the deterministic cointegration does not have severe size distortion problem and has fairly high size adjusted power against no cointegration.

simultaneously (see, e.g., Hansen and Singleton (1982) and Mehra and Prescott (1985)). Several authors have found evidence that time-nonseparability in preferences could help explain asset returns as we discussed in Section II.[13]

The econometric model for our GMM procedure is based on the aggregate version of the asset pricing equation (13), which implies $E_t(\epsilon_g^0(t))=0$, where

$$\epsilon_g^0(t) = \beta[(C_a^n(t+1)+\lambda^n C_a^n(t))^{-\alpha} + \lambda^n\beta(C_a^n(t+2)+\lambda^n C_a^n(t+1))^{-\alpha}]R(t+1)$$

$$-[(C_a^n(t)+\lambda^n C_a^n(t-1))^{-\alpha} + \lambda^n\beta(C_a^n(t+1)+\lambda^n C_a^n(t))^{-\alpha}] \qquad (18),$$

where C_a^n indicates aggregate nondurable consumption. We define $\epsilon_g(t)=\epsilon_g^0(t)/[(1+\lambda^n)C_a^n(t)]^{-\alpha}$, and use $\epsilon_g(t)$ as the disturbance for the GMM estimation. Since $[(1+\lambda^n)C_a^n(t)]^{-\alpha}$ is in the information available at t, $E_t(\epsilon_g(t))=0$. We scale the disturbance by $C_a^{-\alpha}$ to achieve stationarity required for the GMM and $(1+\lambda^n)^{-\alpha}$ to avoid a trivial solution $\beta=0$, $\alpha=\infty$ that causes an identification problem. The disturbance term is MA of order one because of the time-nonseparable 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^n. We use the same weighting matrix to form unrestricted estimates. We then

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[13]We do not take account of the time aggregation problem that Heaton (1991) studied for an asset pricing model with time-nonseparable 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 cover an even longer period, the time aggregation problem for the asset pricing equation can be severe for annual data.

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^n, and α are estimated. The difference is Eichenbaum, Hansen, and Singleton's (1988) C_T test, which has an asymptotic chi-square distribution with one degree of freedom. In our empirical work, we often use three instrumental variables, so that unrestricted GMM estimation is just identified and J_T is zero. In this case, C_T coincides with restricted J_T. 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 $T^{1/2}$.

IV. Data

Quarterly seasonally adjusted data in the National Income and Product Accounts (NIPA) were used for consumption. We used three alternative measures of nondurable consumption: nondurables plus services (NDS), nondurables (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 labor 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 data. Third, the long-run IES for food is a key parameter for Atkeson and Ogaki's (1991) model.

The per capita real consumption series was constructed by dividing the aforementioned series of constant 1982 dollar consumption by the quarterly average of monthly civilian noninstitutional adult (age sixteen 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's contributions to social security, and other labor income (benefits).[14] 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 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.[15] Real gross returns were obtained from the nominal returns and the implicit 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 1989:I.

V. 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

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[14]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.

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

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 of 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 1 reports CCR results.[16] 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

[16]We used Ogaki's 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 (1990) 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 (1990) automatic bandwidth estimator, S_T, was constructed from fitting AR(1) to each disturbance.

positive sign. Our estimates indicate that the long-run IES ($1/\alpha$) is smaller for food, supporting Atkeson and Ogaki's model as we discussed.

Point estimates are strikingly similar when different regressands are used. This is favorable evidence for cointegration because we expect similar estimates when the 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 percent level. We only rejected the null at the 5 percent level when ln(W) was used as the regressand for NDS. The $H(0,1)$ tests the deterministic cointegration restriction. We always fail to reject this restriction at the 5 percent level.

Table 2 presents results for the specification test based on the asset pricing equation. For each of three measures of nondurable consumption, the estimate of α from the cointegrating regression with the log real wage rate as the regressand was used for the GMM procedure.[17] The first panel reports

[17] We used Hansen-Heaton-Ogaki GAUSS GMM Package that was supported by NSF Grant SES-3512371 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 (1989) up to ten 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). The zero restrictions on the covariance matrix from the MA(1) structure of the disturbance term to calculate the weighting matrix were successfully imposed for all final iterations except for NDS and ND when multiple returns were used. For these cases, iterations lead to estimates of λ^n that were close to -1 and the nonlinear search failed as described in the text before the tenth iterations. We report the first

results when only the value weighted returns are used. The instrumental variables used for $\epsilon_g(t)$ were a constant, $C_a^n(t-1)/C_a^n(t-2)$, and a gross real stock return lagged one period. The parameters estimated were β and λ^n. Our estimates of β are less than one and seem reasonable. Our estimates of λ^n are less than one in absolute value and also seem reasonable. We estimate λ^n to be statistically significantly negative, implying habit formation. The specification test, C_T test, does not reject our model at conventional levels.

The second panel of Table 2 reports results when both stock returns and nominal risk free returns are used. In order to keep the number of instrumental variables small as Tauchen (1986) and Kocherlakota (1989) recommend for better small sample properties, we used only a constant for the disturbance for stock returns and a constant and a real gross Treasury bill return lagged one period for the disturbance for Treasury bill returns. When values of λ^n tried in the nonlinear search routine became close to -1, the routine often failed because $C_a^n(t)+\lambda^n C_a^n(t-1)$ was negative for some t. We could not obtain any results for food with this set of instruments because of this problem. Instrumental variables for the results reported in Table 2 for food were a constant, $C_a^n(t-1)/C_a^n(t-2)$, and gross real stock return lagged one period for the disturbance for stock returns and a constant, $C_a^n(t-1)/C_a^n(t-2)$, and a gross real Treasury bill return lagged one period for the disturbance for Treasury bill returns.[18] The C_T test statistic was significant at the five percent level only for ND, where it is

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iteration results for these two cases.

[18] We also tried this set of instruments for NDS and ND. Values tried in the nonlinear search routine for λ^n became close to -1 for unrestricted estimation and we could not obtain unrestricted estimates.

only marginally significant. Point estimates of β are larger than those reported in the first panel but still less than one for NDS and ND. Point estimates for β are slightly larger than one, but are not statistically significantly larger than one. Estimates of λ^n are significantly larger than one at the five percent level and thus reasonable. Except for food, estimates of λ^n are significantly different from zero at the five percent level.[19]

VI. Conclusions

In this paper we tested the following three implications of a neoclassical labor 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 nondurable 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.

We used three alternative measures of nondurable consumption, NDS, ND, and food. We did not reject any of the three implications of our model at the one percent level. Only a few of our specification test statistics were significantly large at the five percent level. Our estimates of the

[19] 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 $C^n_a(t)$ by $C^n_a(t)-\gamma$, where γ is the subsistence level of nondurable consumption. We used Ogaki and Atkeson's (1991b) estimates in a way similar to that of Ogaki and Atkeson (1991a). We found very little effects on our empirical results from fixed subsistence levels.

long-run IES, $1/\alpha$, were about 0.8 for NDS, 0.4 for ND, and 0.3 for food.

The results just described stand in sharp contrast to those usually reported in the literature on the estimation and testing of the neoclassical representative agent model. It is now fairly commonplace to regard such empirical work as a parameter estimation exercise and to not be surprised if the model is rejected formally. Our results suggest that treating the information in real wages and consumption in the way that is implied by the optimal labor contract model yields sensible parameter estimates and does not lead to the rejection of the model.

REFERENCES

Abowd, John M. and David Card. "Intertemporal Labor Supply and Long-Term Employment Contracts." *American Economic Review* 77 (March 1987): 50-68.

Altonji, Joseph, and Ashenfelter, Orley. "Wage Movements and the Labour Market Equilibrium Hypothesis." *Economica* 47 (August 1980): 217-245.

Andrews, Donald K. "Heteroskedasticity and Autocorrelation Consistent Covariance Matrix Estimation." *Econometrica* 59 (May 1991): 817-58.

Andrews, Donald K. and Monahan, J. Christopher. "An Improved Heteroskedasticity and Autocorrelation Consistent Covariance Matrix Estimator." Manuscript. New Haven: Yale University, 1990.

Atkeson, Andrew, and Ogaki, Masao. "Does the Intertemporal Elasticity of Substitution Vary with Wealth and Does That Matter for Aggregate Saving?" Manuscript, 1991.

Azariades, Costas. "Implicit Contracts and Unemployment Equilibria." *Journal of Political Economy* 83 (December 1975): 1183-202.

Benhabib, Jess; Rogerson, Richard; and Wright, Randall. "Homework in Macroeconomics II: Aggregate Fluctuations." Manuscript. (1990)

Blundell, Richard; Browning, Martin; and Meghir, Costas. "A Microeconometric Model of Intertemporal Substitution and Consumer Demand." Manuscript. 1990

Braun, R. Anton. "How Large is the Optimal Inflation Tax?" Manuscript. 1991

Christiano, Lawrence J., and Eichenbaum, Martin. "Unit Roots in Real GNP - Do We Know and Do We Care?" *Carnegie-Rochester Conference Series on Public Policy* 32 (Spring 1990): 7-62.

Cochrane, John H. "How Big is the Random Walk in GNP?" *Journal of Political Economy*, 96 (October 1988): 893-920.

Constantinides, George M. "Habit Formation: A Resolution of the Equity Premium Puzzle." *Journal of Political Economy* 98 (June 1990): 519-43.

Eichenbaum, Martin, and Hansen, Lars Peter. "Estimating Models with Intertemporal Substitution Using Aggregate Time Series Data," *Journal of Business & Economic Statistics*, 8 (January 1990): 53-69.

Eichenbaum, Martin S.; Hansen, Lars Peter; and Singleton, Kenneth J. "A Time Series Analysis of Representative Agent Models of Consumption and Leisure Choice Under Uncertainty," *Quarterly Journal of Economics*, 103 (February 1988) 51-78.

Engle, Robert F. and Granger, C.W.J. "Co-Integration and Error Correction: Representation, Estimation, and Testing," *Econometrica*, 55, (March 1987): 251-276.

Ferson, Wayne E. and Constantinides, George M. "Habit Formation and Durability in Aggregate Consumption: Empirical Tests." *Jouranl of Financial Economics* 29, forthcoming (1991).

Gonzalo, Jesús. "Cointegration and Aggregation." UCSD Discussion paper 89-30. San Diego: University of California, San Diego. (1989).

Greenwood, Jeremy and Hercowitz, Zvi. "The Allocation of Goods and Time over the Business Cycle." Federal Reserve Bank of Minneapolis Discussion Paper 26. (1990).

Hall, Robert E. "Stochastic Implications of the Life Cycle-Permanent Income Hypothesis: Theory and Evidence." *Journal of Political Economy* 86 (December 1978): 971-87.

Han, Hsiang-Ling, and Ogaki, Masao. "Consumption, Income, and Cointegration: Further Analysis. Manuscript in progress (1991).

Hansen, Bruce E. "Efficient Estimation of Cointegrating Vectors in the Presence of Deterministic Trends," manuscript, University of Rochester, 1990.

Hansen, Lars Peter. "Large Sample Properties of Generalized Method of Moments Estimators," *Econometrica* 50 (July 1982): 1029-1054.

Hansen, Lars Peter, and Singleton, Kenneth J. "Generalized Instrumental Variable Estimation of Nonlinear Rational Expectations Models," *Econometrica*, 50 (September 1982): 1269-1286.

Hausman, J. A. "Specification Tests in Econometrics." *Econometrica* 46 (November 1978): 1251-71.

Heaton, John. "The Interaction between Time-Nonseparable Preferences and Time Aggregation." Working Paper #3181-90-EFA, Sloan School of Management, Massachusetts Institute of Technology, 1990.

———. "An Empirical Investigation of Asset Pricing with Temporally Dependent Preference Specification." Working Paper #3245-91-EFA, Sloan School of Management, Massachusetts Institute of Technology, 1991.

Hotz, V. Joseph; Kydland, Finn E.; and Sedlacek, Bulherme L. "Intertemporal Preferences and Labor Supply." *Econometrica* 56 (March 1988): 335-360.

Houthakker, H.S. "Additive Preferences." *Econometrica* 28 (April 1960): 244-56.

Johansen, Søren, 1988, Statistical analysis of cointegration vectors, *Journal of Economic Dynamics and Control* 12 (1988), 231-254.

Kennan, John. "An Econometric Analysis of Fluctuations in Aggregate Labor Supply and Demand." *Econometrica* 56 (March 1988): 317-333.

King, Robert G., Charles I. Plosser, and Sergio T. Rebelo. "Production, Growth, and Business Cycles: I. The Basic Neoclassical Model." *Journal of Monetary Economics* 21 (March/May 1988): 195-232.

Kocherlakota, Narayana R. "On Tests of the Time and State Separable Utility-Based Asset Pricing Model." Manuscript, (October 1989).

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

Kydland, Finn E. "Labor-Force Heterogeneity and the Business Cycle." *Carnegie-Rochester Conference Series on Public Policy* 21 (1984): 173-208.

Kydland, Finn E. and Prescott, Edward C. "Time to Build and Aggregate Fluctuations." *Econometrica* 50 (November 1982): 1345-70.

Mankiw, N. Gregory; Rotemberg, Julio J.; and Summers, Lawrence H. "Intertemporal Substitution in Macroeconomics," *Quarterly Journal of Economics*, C (1985): 225-52.

Mehra, Rajnish and Prescott, Edward C. "The Equity Premium: A Puzzle." *J.Monetary Econ*. 15 (January 1984): 145-61.

Nelson, Charles R. and Plosser, Charles I. "Trends and Random Walks in Macroeconomic Time Series." *J. Monetary Econ*. 10 (September 1982): 139-62.

Ogaki, Masao. "Aggregation of Intratemporal Preferences Under Complete Markets." Rochester Center for Economic Research Working Paper No. 249, University of Rochester (September 1990).

———. "Growth in Open Economies: A Comment." Forthcoming in *Carnegie-Rochester Conference Series on Public Policy*, 1991.

Ogaki, Masao, and Andrew Atkeson. "Estimating a Model of Engel's Law, Saving, and Risk Bearing with Indian Panel and U.S. Time Series Data." Manuscript, 1991. (a)

———. "Subsistence Levels, Risk Sharing, and Saving: Evidence from Indian Panel Data." Manuscript, 1991. (b)

Ogaki, Masao, and Park, Joon Y. "A Cointegration Approach to Estimating Preference Parameters." Paper presented at the Conference on Empirical Applications of Structural Models organized by *Econometrica* and Social Systems Research Institute, University of Wisconsin, Madison, Wisconsin, (May 1990).

Park, Joon Y. "Testing for Unit Roots and Cointegration by Variable Addition." *Advances in Econometrics*, 8, (1990): 107-33.

———. "Canonical Cointegrating Regressions," CAE Working Paper No.88-29R, forthcoming in *Econometrica* (1990).

Park, Joon Y. and Choi, Buhmsoo. "A New Approach to Testing for a Unit Root," CAE Working Paper No. 88-23, Cornell University, 1988.

Park, Joon Y., Ogaki, Masao. "Inference in Cointegrated Models Using VAR Prewhitening to Estimate Shortrun Dynamics." Rochester Center for Eocnomic Research Working Paper No. 281, University of Rochester, 1991.

Park, Joon Y.; Ouliaris Sam, and Buhsoo Choi. "Spurious Regressions and Tests for Cointegration," CAE Working Paper No. 88-07, Cornell University, 1988.

Phillips, P.C.B. "Spectral regression for cointegrated time series, Cowles Foundation Discussion Paper no. 872., Yale University, New Haven, CT, 1988.

———. "Optimal Inferences in Cointegrated Systems, Econometrica, 59 (March 1991): 283-306.

Phillips, Peter C. B., and Hansen, Bruce E. "Statistical Inference in Instrumental Variables Regression with I(1) Processes," *Review of Economic Studies* 57 (January 1990): 99-125.

Phillips, Peter C. B. and Perron, Pierre. "Testing for a Unit Root in Time Series Regression," *Biometrica*, 75 (1988): 335-46.

Rebelo, Sergio. "Growth in Open Economies." forthcoming in *Carnegie-Rochester Conference on Public Policy*, 1991.

Rosen, Sherwin. "Implicit Contracts: A Survey." *J.Economic Literature* 23 (September 1985): 1144-75.

Said, Said D. and Dickey, David A. "Testing for Unit Roots in Autoregressive-Moving Average Models of Unknown Order." *Biometrica* 71 (1984): 599-607

Saikkonen, P.Asymptotic efficient estimation of cointegration regressions, Mimeo. University of Helsinki, Helsinki, Finland, 1989.

Stock, James H., and Watson, Mark W. "A Simple MLE of Cointegrating Vectors in Higher Order Integrated Systems," NBER Technical Working Paper No. 83, (December 1989).

Tauchen, George. "Statistical Properties of Generalized-Method-of-Moments Estimators of Structural Parameters Obtained from Financial Market Data," *Journal of Business and Economic Statistics* 4 (October 1986): 397-425.

West, Kenneth D. "Asymptotic Normality, When Regressors Have a Unit Root," *Econometrica* 56 (November 1988): 1397-1417.

Wright, Randall D. "The Observational Implications of Labor Contracts in a Dynamic General Equilibrium Model." *Journal of Labor Economics* 6 (October 1988): 530-51.

TABLE 1

CANONICAL COINTEGRATING REGRESSION RESULTS

Regressand[a]	α [b]	$1/\alpha$ [b]	$H(0,1)$[c]	$H(1,2)$[d]	$H(1,3)$[d]	$H(1,4)$[d]
w, NDS	1.224	0.817	3.562	0.838	6.738	10.731
	(0.123)	. . .	(0.059)	(0.360)	(0.034)	(0.013)
c, NDS	1.309	0.764	0.197	0.038	2.613	3.076
	. . .	(0.037)	(0.657)	(0.846)	(0.271)	(0.380)
w, ND	2.442	0.413	1.166	0.020	4.448	8.090
	(0.115)	. . .	(0.280)	(0.886)	(0.109)	(0.044)
c, ND	2.439	0.410	1.389	0.001	0.113	0.353
	. . .	(0.027)	(0.239)	(0.980)	(0.945)	(0.950)
w, food	3.114	0.321	1.430	0.413	0.468	1.784
	(0.275)	. . .	(0.232)	(0.520)	(0.791)	(0.618)
c, food	3.027	0.330	0.667	0.028	0.646	2.245
	. . .	(0.031)	(0.414)	(0.868)	(0.724)	(0.523)

NOTE: Park and Ogaki's (1991) method with Andrews's (1991) automatic bandwidth parameter estimator was used to estimate long-run correlation parameters.

[a]In this column, 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.

[b]Standard errors are in parentheses.

[c]P-values are in parentheses. This statistic tests the deterministic cointegration restriction.

[d]P-values are in parentheses. These tests the null of stochastic cointegration.

TABLE 2

Specification Test Results Based on the Asset Pricing Equation

Consumption[a]	α^{b}	β^{b}	λ^{nb}	J_T^{c}	df	C_T^{d}
			Stock Returns			
NDS (R)	1.224	0.982 (0.006)	-0.530 (0.082)	. . .		1.829 (0.176)
ND (R)	2.422	0.983 (0.006)	-0.279 (0.096)	. . .		0.193 (0.660)
Food (R)	3.114	0.982 (0.007)	-0.267 (0.044)	. . .		1.074 (0.300)
			Stock Returns and Treasury bill Returns			
NDS (R)	1.224	0.999 (0.004)	-0.706 (0.094)	. . .		3.219 (0.073)
ND (R)	2.422	0.998 (0.003)	-0.454 (0.090)	. . .		4.069 (0.044)
Food (R)	3.114	1.001 (0.002)	-0.105 (0.065)	7.279 (0.122)	4	. . .
Food (U)	3.849 (1.244)	1.001 (0.002)	-0.079 (0.064)	7.011 (0.072)	3	0.268 (0.605)

NOTE: Our estimate of α reported Table 1 with ln(W) as the regressand was used for each consumption category for restricted estimation.

[a]In this column, we report the measure of nondurable consumption used and whether or not the estimation is restricted (indicated by R) or unrestricted (indicated by U).

[b]Standard errors are in parentheses when unrestricted.

[c]This statistic tests the overidentifying restrictions. This is reported only it differs from the C_T statistic. The p-values are in parentheses.

[d]This chi-square statistic with one degree of freedom tests the restriction imposed. The p-values are in parentheses.