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An Empirical Assessment of a Consumption CAPM with a Reference Level under Incomplete Consumption Insurance

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Abstract

We study asset pricing implications of the preference specification in which an agent derives utility from both the ratio of his consumption to some reference level and this level itself under incomplete consumption insurance and limited asset market participation. Assuming that the reference level responds gradually to changes in aggregate consumption per capita, we show that when asymmetry in individual consumption is taken into account, the obtained estimate of the elasticity of intertemporal substitution is in the conventional range and significantly different from the inverse of the relative risk aversion (RRA) coefficient as the definition of assetholders is tightened. Both the power utility model and the ratio preference specification are rejected statistically.

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1 Introduction

The standard consumption capital asset pricing model treats asset prices as being determined by the consumption and savings decisions of a single representative agent assumed to have the conventional time- and state-separable power utility specification. Empirical tests of the representative-agent model reject the model in several ways. Thus, Mehra and Prescott (1985) show that the representative-agent model is not able to explain the observed average excess return on the stock market unless risk aversion is assumed to be implausibly high (the equity premium puzzle). The large estimate of risk aversion implies another puzzle: if investors are extremely risk-averse, then the observed average growth rate of per capita consumption is consistent with the low short-term real interest rate only if the representative agent has a negative rate of time preference. This is the risk-free rate puzzle (Weil (1989)). Ferson and Constantinides (1991), Hansen and Jagannathan (1991), and Hansen and Singleton (1982, 1983) test the conditional Euler equations for an assumed representative agent and find that the overidentifying restrictions strongly reject the model when the Euler equations for the equity premium and the risk-free rate are estimated jointly.

Since Mehra and Prescott's original investigation, several generalizations of essential features of the representative-agent model have been suggested to mitigate its poor empirical performance. Thus, Brav, Constantinides, and Géczy (2002), Constantinides and Duffie (1996), Mankiw (1986), Mehra and Prescott (1985), Semenov (2002), and Weil (1992) suggest that deviations from complete consumption insurance have the potential to explain the equity premium and risk-free rate puzzles. In particular, Semenov (2002) develop an approximate equilibrium multifactor model for expected returns in which the priced risk factors are cross-moments of return with the moments of individual consumption and find that the model can explain both the equity premium and the risk-free rate with economically plausible (less than 2) values of risk aversion and the time discount factor when the agent's marginal utility of consumption is expanded as a Taylor series up to terms capturing the skewness of the distribution of individual consumption around its conditional expectation.

Another possible explanation of empirical rejections of the representative-agent model is excessive rigidity of the conventional time- and state-separable utility function which constrains the elasticity of intertemporal substitution to be the reciprocal of the RRA coefficient. Constantinides (1990) studies an internal habit model in which the utility is a power of the difference between the current consumption flow and a fraction of a weighted sum of lagged consumption flows and proves that habit persistence and/or durability of consumption drives a wedge between the elasticity of consumption with respect to investment returns and the inverse of the RRA coefficient. Epstein and Zin (1989, 1991), hereafter EZ, assume that the agent's lifetime utility depends on both current consumption and a certainty equivalent of a random future utility through an intertemporal constant elasticity of substitution utility function. For the certainty equivalent, EZ (1989, 1991) consider a constant relative risk aversion expected utility specification. This generalized specification of intertemporal utility allows a separation of risk aversion (reflected in the certainty equivalent function) from intertemporal substitution (encoded in the aggregator function).

Garcia, Renault, Semenov (2003a, 2003b), henceforth GRS, propose another way to disentangle intertemporal substitution and risk aversion. They assume an agent to derive utility from both the ratio of his consumption to some benchmark level of consumption and this level itself. They show that if the external reference level matters for a decision maker and the reference consumption level growth rate is correlated with the market portfolio return, this expected utility model has the ability to explain both the equity premium and the risk-free rate as well as to separate the elasticity of intertemporal substitution from the inverse of the RRA coefficient. An important result is that if the reference consumption level growth rate is assumed to be a function of the market portfolio return alone, this utility specification yields a SDF which is isomorphic in its pricing implications to the pricing kernel corresponding to the EZ (1989, 1991) non-expected utility specification. The comparison between the EZ (1989, 1991) non-expected utility model and the GRS (2003a) expected utility model with a reference level shows that the elasticity of intertemporal substitution remains the same in the two models, while the measure of risk aversion in the GRS (2003a) utility specification differs from that in the EZ (1989, 1991) utility model. An attractive feature of the GRS (2003a) preference specification is that, in contrast to the Constantinides (1990) internal habit model, consumption is not required to be always above the reference level for marginal utility to be positive. GRS (2003a) test this utility function under the assumption of complete consumption insurance and obtain the point estimate of the elasticity of intertemporal substitution that is in the conventional range and statistically different from the inverse of the RRA coefficient. Besides, their empirical result is that the SDF corresponding to the preference specification with a reference level outperforms the EZ (1989, 1991) pricing kernel.

The goal of this paper is to examine the asset pricing implications of the preference specification with a reference level under the assumptions of incomplete consumption insurance and limited participation of consumers in the asset markets using the approximate equilibrium multifactor model for expected asset returns developed in Semenov (2002). The common to all agents contemporaneous macroeconomic factors posited to affect the reference level are assumed to be adequately proxied by the level of aggregate consumption per capita. Assuming further the substistence level to response gradually to changes in aggregate consumption per capita, we use the following two-stage procedure to estimate the parameters of interest. In the first step, we estimate sensitivity of the reference level to changes in aggregate consumption per capita. The second step is to use the iterated generalized method of moments (GMM) approach to estimate the conditional Euler equations for the equity premium and the risk-free rate of return implied by the Semenov (2002) approximate equilibrium multifactor model for expected asset returns and the GRS (2003a) preference specification using the estimate of the speed of adjustment parameter obtained in the first step.

The remaining of this paper is organized as follows. In Section 2, we briefly review the major features of the approximate equilibrium multifactor model for expected asset returns developed in Semenov (2002). Section 3 details the preference specification with a reference level responding gradually to changes in aggregate consumption per capita. Section 4 describes the data, estimation and testing methodology and presents estimation results. Section 5 concludes.

2 The Equilibrium Multifactor Pricing Model

Consider the intertemporal consumption and portfolio choice problem of a single representative investor who can trade freely in asset i and who maximizes expected lifetime discounted utility

$$MaxE_t\left[\sum_{j=0}^{\infty}\delta^j u\left(C_{k,t+j}\right)\right] \tag{1}$$

subject to his budget constraint

$$C_{k,t+1} - R_{i,t+1}(W_{k,t} - C_{k,t}) = 0, \tag{2}$$

where δ is the subjective discount factor, $C_{k,t+j}$ is the investor's consumption in period $t+j, u\left(C_{k,t+j}\right)$ is the one-period utility of consumption at $t+j$, $W_{k,t}$ is the investor's welfare in period t, $R_{i,t+1}$ is the simple gross return on asset i, and $E_t\left[\cdot\right]$ denotes the mathematical expectation conditioned on the period-t information set, Ω_t, that is common to all agents.[1]

The first-order condition describing the investor's optimal consumption and portfolio plan is

$$\delta E_t\left[u'\left(C_{k,t+1}\right)R_{i,t+1}\right] = u'\left(C_{k,t}\right),\ k=1,...,K,\ i=1,...,I. \tag{3}$$

The right side of (3) is the loss in utility if the investor buys another unit of the asset, the left side is the increase in discounted, expected utility he obtains from the extra payoff at

[1] $u\left(\cdot\right)$ is assumed to be increasing, strictly concave, and differentiable.

time $t+1$. Hence, in the optimum the investor equates the marginal loss and the marginal gain from holding of his portfolio.

Assuming $u(\cdot)$ to be $N+1$ times differentiable, Semenov (2002) uses a N-order Taylor expansion to the individual k's marginal utility around the conditional expectation of consumption, $h_{t+1} \equiv E_t[C_{k,t+1}]$:[2]

$$u'(C_{k,t}) = \sum_{n=0}^{N} \frac{1}{n!} u^{(n+1)}(h_t)(C_{k,t} - h_t)^n, \; k = 1, ..., K. \tag{4}$$

Substituting for $u'(C_{k,t})$ from (4), we obtain

$$\delta \sum_{n=0}^{N} \frac{1}{n!} u^{(n+1)}(h_{t+1}) E_t[(C_{k,t+1} - h_{t+1})^n R_{i,t+1}] = \sum_{n=0}^{N} \frac{1}{n!} u^{(n+1)}(h_t)(C_{k,t} - h_t)^n, \tag{5}$$

$k = 1, ..., K$, $i = 1, ..., I$.

By summing these equations over individuals and dividing by the number of individuals in the population, we obtain the following set of equations:

$$\delta \sum_{n=0}^{N} \frac{1}{n!} u^{(n+1)}(h_{t+1}) E_t[Z_{n,t+1} R_{i,t+1}] = \sum_{n=0}^{N} \frac{1}{n!} u^{(n+1)}(h_t) Z_{n,t}, \tag{6}$$

$i = 1, ..., I$, where $Z_{n,t} \equiv \frac{1}{K} \sum_{k=1}^{K} (C_{k,t} - h_t)^n$.

Equation (6) can be rewritten as

$$E_t[R_{i,t+1}] = \delta^{-1} \sum_{n=0}^{N} \frac{1}{n!} \frac{u^{(n+1)}(h_t)}{u'(h_{t+1})} Z_{n,t} - \sum_{n=1}^{N} \frac{1}{n!} \frac{u^{(n+1)}(h_{t+1})}{u'(h_{t+1})} \cdot E_t[Z_{n,t+1} R_{i,t+1}]. \tag{7}$$

This is the approximate equilibrium multifactor model for expected asset returns.[3]

For the expected excess return on the market portfolio over the risk-free rate, $RP_{t+1} \equiv R_{M,t+1} - R_{f,t+1}$, equation (6) reduces to

$$\sum_{n=0}^{N} \frac{1}{n!} u^{(n+1)}(h_{t+1}) E_t[Z_{n,t+1} RP_{t+1}] = 0. \tag{8}$$

The Euler equation for the expected equilibrium risk-free rate is

$$\delta \sum_{n=0}^{N} \frac{1}{n!} u^{(n+1)}(h_{t+1}) E_t[Z_{n,t+1} R_{f,t+1}] = \sum_{n=0}^{N} \frac{1}{n!} u^{(n+1)}(h_t) Z_{n,t}. \tag{9}$$

[2]Here, and throughout the paper, $u^{(n)}(\cdot)$ denotes the nth derivative of $u(\cdot)$.
[3]See Semenov (2002).

3 Preferences

Following GRS (2003a), consider a single investor whose period t utility function is given by

$$u\left(C_{k,t}, S_{k,t}\right) = \frac{\left(\frac{C_{k,t}}{S_{k,t}}\right)^{1-\gamma} S_{k,t}^{1-\varphi} - 1}{1-\gamma}. \tag{10}$$

Here, $S_{k,t}$ is the agent's k time-varying subsistence or reference consumption level in period t, γ is the Arrow-Pratt measure of relative risk aversion, and the parameter φ controls the curvature of utility over the reference level of consumption.

Utility function (10) nests some preference specifications which can be obtained given different values of the curvature parameter φ. Thus, if $\varphi = \gamma$, the reference consumption level plays no role in asset pricing and we get the standard time-separable power utility model. When $\varphi < \gamma$, an increase in the reference level raises the marginal utility of the agent's own consumption. Gali (1994) refers to this type of externalities as positive consumption externalities. Alternatively, when $\varphi > \gamma$, an increase in the benchmark level lowers the marginal utility of consumption. These are negative consumption externalities.[4] With $\varphi = 1$, we obtain the ratio preference specification when the agent derives utility from consumption relative to the benchmark level. If $\varphi \neq \gamma$ and $\varphi \neq 1$, then the agent takes into account both the ratio of his consumption to the subsistence level and this level itself when choosing how much to consume.

Assume the time-varying substistence level to be unaffected by any one agent's consumption decisions. GRS (2003a) show that if the reference consumption level is exogenous to an individual consumer, this utility specification not only has the potential to explain the equity premium and risk-free rate puzzles but also allows to disentangle intertemporal sub-

[4] In the special case, when the reference level is proxied by past consumption levels, positive consumption externalities are usually referred to as habit persistence in preferences, while negative consumption externalities correspond to durability in consumption expenditures (see, for example, Eichenbaum and Hansen (1990), Eichenbaum, Hansen, and Singleton (1988), Ferson and Constantinides (1991), Gallant and Tauchen (1989), and Heaton (1995)).

stitution and risk aversion.[5] In particular, they show that in this model the intertemporal elasticity of consumption is

$$\sigma_k = \frac{\partial \triangle c_{k,t+1}}{\partial r_{i,t+1}} = \frac{1 + (\gamma - \varphi) \frac{\partial \triangle s_{k,t+1}}{\partial r_{i,t+1}}}{\gamma}, \tag{15}$$

where $\frac{\partial \triangle s_{k,t+1}}{\partial r_{i,t+1}}$ can be interpreted as the elasticity of the subsistence level with respect to investment returns. Equation (15) implies that the elasticity of intertemporal substitution differs from the inverse of the RRA coefficient if $(\gamma - \varphi) \frac{\partial \triangle s_{k,t+1}}{\partial r_{i,t+1}} \neq 0$.[6][7]

Since the reference level of consumption is not observable, this model is of little use without specifying a way to measure the factors that are posited to affect subsistence requirements. In GRS (2003a), it is shown that, given different assumptions about the reference level generating process, the pricing kernel corresponding to preference specification (10) nests several the most often used in asset pricing stochastic discount factors (SDFs). Thus, when the benchmark level of consumption is assumed to be determined by past consumption levels only, the model generalizes the usual external habit formation specifications. One

[5]Since the reference consumption level is external, the SDF is then

$$M_{t+1} \equiv \delta \frac{u'(C_{k,t+1}, S_{k,t+1})}{u'(C_{k,t}, S_{k,t})} = \delta \left(\frac{C_{k,t+1}}{C_{k,t}} \right)^{-\gamma} \left(\frac{S_{k,t+1}}{S_{k,t}} \right)^{\gamma - \varphi}. \tag{11}$$

Under the assumption that there is an representative agent, we can rewrite (11) as

$$M_{t+1} = \delta \left(\frac{C_{t+1}}{C_t} \right)^{-\gamma} \left(\frac{S_{t+1}}{S_t} \right)^{\gamma - \varphi}. \tag{12}$$

Assuming joint conditional lognormality and homoskedasticity of the consumption growth rate and asset returns, GRS (2003a) obtain

$$r_{f,t+1} = -log\delta + \gamma E_t [\Delta c_{t+1}] - \frac{1}{2}\gamma^2 \sigma_c^2 - (\gamma - \varphi) E_t [\Delta s_{t+1}] - \frac{1}{2} (\gamma - \varphi)^2 \sigma_s^2 + \gamma (\gamma - \varphi) \sigma_{cs} \tag{13}$$

and

$$E_t [r_{i,t+1} - r_{f,t+1}] = -\frac{1}{2}\sigma_i^2 + \gamma \sigma_{ic} - (\gamma - \varphi) \sigma_{is}, \tag{14}$$

where Δc_{t+1} is the log of the consumption growth rate, Δs_{t+1} is the log of the reference consumption level growth rate, $r_{i,t+1}$ is the log of the simple gross return on asset i, and σ_{xy} denotes the unconditional covariance of innovations. The first three terms on the right-hand side of (13) and the first two terms on the right-hand side of (14) are the same as for a time-separable power utility function of consumption alone. Thus, utility function (10) has the ability to explain the equity premium puzzle if the term $-(\gamma - \varphi)\sigma_{is}$ is positive and the risk-free rate puzzle if the term $-(\gamma - \varphi) E_t [\Delta s_{t+1}] - \frac{1}{2} (\gamma - \varphi)^2 \sigma_s^2 + \gamma (\gamma - \varphi) \sigma_{cs}$ is negative.

[6]Since $\frac{\partial \triangle s_{t+1}}{\partial r_{i,t+1}}$ and σ_{is} have the same sign, if utility specification (10) contributes towards a solution of the equity premium puzzle, it also yields the elasticity of intertemporal substitution which is less than the inverse of the risk aversion coefficient (see equations (14) and (15)).

[7]Another example of the utility specification allowing to separate the elasticity of intertemporal substitution from risk aversion in the expected utility framework is the Ferson-Constantinides (1991) internal habit model, in which the utility is a power function of the difference between the current consumption flow and a fraction of a weighted sum of lagged consumption flows. However, this model is restrictive in that consumption must always be above habit for marginal utility to be positive, what is not required in model (10).

reasonable approach is to assert that the agent's reference level could be affected not only by past consumption, but also by some contemporaneous macro- and microeconomic factors such as business cycles, inflation, age of reference person, his education, marital status, etc. GRS (2003a, 2003b) demonstrate that if we assume that the return on the market portfolio is a valid proxy for the common to all agents macroeconomic factors and the benchmark consumption level does not depend on past consumption, preference specification (10) yields a SDF which is observationally equivalent to the pricing kernel corresponding to the EZ (1989, 1991) non-expected recursive utility function.[8]

In this paper, we assume that the common to all agents contemporaneous macroeconomic factors posited to affect the reference level may be adequately proxied by the level of aggregate consumption per capita. Assume further that the substistence level responses gradually to changes in aggregate consumption per capita and the dynamics of $\{logS_{k,t+1}\}$ are given by the following equation:

$$logS_{k,t+1} = a_{k,t+1} + (1 - \lambda_k)\, logS_{k,t} + \lambda_k logC_{t+1},\ 0 \leqslant \lambda_k \leqslant 1, \tag{16}$$

where $a_{k,t+1}$ is the rate of reference level growth caused by the increase in the standard of living.[9]

If we repeatedly lag and substitute equation (16), we can write $logS_{k,t+1}$ as a weighted sum of current and past measured consumption:

$$logS_{k,t+1} = \frac{a_{k,t+1}}{\lambda_k} + \lambda_k \sum_{i=0}^{\infty} (1 - \lambda_k)^i\, logC_{t+1-i},\ 0 < \lambda_k \leqslant 1. \tag{17}$$

Let us assume the reference consumption level to be the same for all agents, $S_{k,t+1} = S_{t+1}$ for all k ($\lambda_k = \lambda$ and $a_{k,t+1} = a_{t+1}$ for all k). Assume further that the rate of growth of the reference consumption level with the passage of time, a_{t+1}, is correlated with the return on the market portfolio, $a_{t+1} = a + b \cdot r_{M,t+1}$, where $r_{M,t+1}$ is the continuously compounded market portfolio return.

[8]Empirical evidence in GRS (2003a) is that when the representative agent's reference consumption level is assumed to depend on both the market portfolio return and lagged aggregate consumption per capita, we are able to fit empirical data on asset returns with economically plausible and statistically significant values of risk aversion and the elasticity of intertemporal substitution (the null hypothesis the elasticity of intertemporal substitution equals the inverse of the RRA coefficient is rejected statistically at the 5% significance level), in opposite to the Epstein-Zin (1989, 1991) SDF which yields a negative estimate of elasticity of substitution.

[9]The higher the value of λ_k, the more rapid the adjustment process. If $\lambda_k = 0$, then $S_{k,t+1} = exp\,(a_{k,t+1}) \cdot S_{k,t}$ at all t (the reference level grows simply with the passage of time). At the other extreme, if $\lambda_k = 1$, there is full adjustment in one period, $S_{k,t+1} = exp\,(a_{k,t+1}) \cdot C_{t+1}$. This case corresponds to the formulation of the benchmark level in Gali (1994) according to which the reference level of consumption only depends on the contemporaneous per capita consumption level in the economy. A similar approach to make the reference level of consumption grow with the passage of time is used in Abel (1999). Specifically, Abel (1999) states $S_t \equiv C_t^{\alpha_0} C_{t-1}^{\alpha_1} \left(G^t\right)^{\alpha_2}$, where G is a constant, $G \geqslant 1$.

When $\lambda = 0$, (16) implies $logS_{t+1} = a + b \cdot r_{M,t+1} + logS_t$. Consequently, $\triangle s_{t+1} = a + b \cdot r_{M,t+1}$ and, therefore, $\frac{\partial \triangle s_{t+1}}{\partial r_{M,t+1}} = b$. From (15), we obtain that all investors have the same elasticity of intertemporal substitution $\sigma = \frac{1+(\gamma-\varphi)b}{\gamma}$. When $\varphi = \gamma$, we get the conventional power utility model for which $\sigma = \frac{1}{\gamma}$ whatever the value of b.

Under the assumptions above, when there is an representative agent, the SDF for model (10) is

$$M_{t+1} = \delta^* \left(\frac{C_{t+1}}{C_t} \right)^{-\gamma} (1 + R_{M,t+1})^{\sigma\gamma-1}, \tag{18}$$

where $\delta^* \equiv \delta \cdot exp\left(a\left(\gamma - \varphi\right)\right)$. This SDF is observationally equivalent to the EZ (1989, 1991) pricing kernel.[10]

If $\lambda > 0$, according to (17)

$$\triangle s_{t+1} = \frac{b}{\lambda}\left(r_{M,t+1} - r_{M,t}\right) + \lambda \triangle c_{t+1} + \lambda \sum_{i=1}^{\infty} (1-\lambda)^i \triangle c_{t+1-i}. \tag{19}$$

It implies $\frac{\partial \triangle s_{t+1}}{\partial r_{M,t+1}} = \frac{b}{\lambda} + \lambda\sigma$ and, therefore, $\sigma = \frac{1+(\gamma-\varphi)\left(\frac{b}{\lambda}+\lambda\sigma\right)}{\gamma}$. This relationship can be rearranged so that we obtain $\sigma = \frac{1+(\gamma-\varphi)\frac{b}{\lambda}}{(1-\lambda)\gamma+\lambda\varphi}$. When $\varphi = \gamma$, our utility specification reduces to the standard time- and state-separable power utility function with $\sigma = \frac{1}{\gamma}$ for any values of λ $(0 < \lambda \leqslant 1)$ and b.

In this paper, we will consider only the case when the rate of reference level growth caused by the increase in the standard of living is constant over time, $b = 0$. We further hypothesize that C_{t+1} is related to S_{t+1} by the relationship $C_{t+1} = S_{t+1} \cdot \varepsilon_{t+1}$, where a disturbance term ε_{t+1} is independent of S_{t+1} and lognormally distributed, $log\varepsilon_{t+1} \sim N\left(0, \sigma_\varepsilon^2\right)$.[11]

Under these assumptions

$$logC_{t+1} = \frac{a}{\lambda} + \lambda \sum_{i=0}^{\infty} (1-\lambda)^i \, logC_{t+1-i} + log\varepsilon_{t+1}, \; 0 < \lambda \leqslant 1. \tag{20}$$

[10] See GRS (2003b) for the detailed comparative analysis of these two SDFs.

[11] A disturbance term ε_{t+1} is assumed to represent a contemporaneous shock to realized aggregate consumption. If the shock is positive, $log\varepsilon_{t+1} > 0$ $(\varepsilon_{t+1} > 1)$, consumption is above the benchmark level. However, when the shock is negative, $log\varepsilon_{t+1} < 0$ $(\varepsilon_{t+1} < 1)$, consumption is presumed to be below the reference level. The only case, when consumption coincides with the benchmark level is the absence of any shock. Under the assumptions $a_{k,t+1} = a$ and $\lambda_k = \lambda$ for all k, substituting $logC_{t+1} = logS_{t+1} + log\varepsilon_{t+1}$ into (16) yields

$$\triangle s_{t+1} = \frac{a}{1-\lambda} + \frac{\lambda}{1-\lambda} log\varepsilon_{t+1}, \; 0 \leqslant \lambda < 1.$$

If $log\varepsilon_{t+1}$ are independent, then $\triangle s_{t+1}$ are IID normal variates with mean $\frac{a}{1-\lambda}$ and variance $\left(\frac{\lambda}{1-\lambda}\right)^2 \sigma_\varepsilon^2$. As λ approaches 1, $var(\triangle s_{t+1})$ approaches $var(\triangle c_{t+1})$.

Equation (20) may be rewritten as

$$logC_{t+1} = \frac{a}{\lambda\left(1-\lambda\right)} + \lambda \sum_{i=1}^{\infty} \left(1-\lambda\right)^{i-1} logC_{t+1-i} + \frac{log\varepsilon_{t+1}}{\left(1-\lambda\right)},\ 0<\lambda<1. \tag{21}$$

Using the Koyck transformation yields

$$\Delta c_{t+1} = \frac{a}{1-\lambda} - \eta_{t+1} + \left(1-\lambda\right)\eta_t,\ 0<\lambda<1, \tag{22}$$

where $\eta_{t+1} \equiv -\frac{log\varepsilon_{t+1}}{(1-\lambda)}$. This is an MA(1) model in which the coefficient of η_t characterizes persistence in the reference consumption level process. This model can be estimated from the time series of aggregate consumption per capita.

In opposite to Cecchetti, Lam, and Mark (1990, 1993) and Kandel and Stambaugh (1991), who assume a persistent discrete-state Markov process for expected aggregate consumption growth z_t, Campbell (1999) assumes z_t to follow an AR(1) process with mean g and persistence ψ:

$$\begin{aligned} \Delta c_{t+1} &= z_t + v_{t+1}, \\ z_{t+1} &= \left(1-\psi\right)g + \psi z_t + u_{t+1},\ \psi>0, \end{aligned} \tag{23}$$

$$\begin{bmatrix} v_{t+1} \\ u_{t+1} \end{bmatrix} \sim N\left(0, \begin{bmatrix} \sigma_v^2 & \rho\sigma_v\sigma_u \\ \rho\sigma_v\sigma_u & \sigma_u^2 \end{bmatrix}\right). \tag{24}$$

Wachter (2002) shows that conditional on consumption data, system (23) has the same likelihood function as the following ARMA(1,1) process:

$$\Delta c_{t+1} = \left(1-\psi\right)g + \psi\Delta c_t + \eta_{t+1} + \theta\eta_t, \tag{25}$$

where

$$\left(\left(\theta+\psi\right)^2 + 1 - \psi^2\right)\sigma_\eta^2 = \sigma_v^2\left(1-\psi^2\right) + \sigma_u^2 \tag{26}$$

and

$$\theta\sigma_\eta^2 = \rho\sigma_v\sigma_u - \psi\sigma_v^2. \tag{27}$$

If $\theta = 0$ ($u_{t+1} = \psi v_{t+1}$), we obtain a linear version of the Mehra-Prescott (1985) model in which consumption growth follows an AR(1) process. Setting $\theta = 0$ and $\psi = 0$ results in the random walk model of consumption (Campbell and Cochrane (1999)). Equation (22) can be obtained from (25) when ψ is set to 0 with the free parameter θ and, hence, is less restrictive than the model assumed by Campbell and Cochrane (1999).

The more general case is to assume that not only the current period shock but also some previous period shocks may affect the level of current consumption: $C_{t+1} = S_{t+1} \cdot \prod_{i=0}^{\infty} \varepsilon_{t+1-i}^{\alpha_i}$, $\alpha_0 = 1$. Under this assumption,

$$\Delta c_{t+1} = \frac{a}{1-\lambda} - \eta_{t+1} - \sum_{i=1}^{\infty} \left(\alpha_i - \alpha_{i-1}\left(1-\lambda\right)\right) \eta_{t+1-i}, \; 0 < \lambda < 1. \tag{28}$$

Another way to take into account some persistence in shocks is to assume an AR(p) model for $log\varepsilon_{t+1}$. If we assume, for example, that $log\varepsilon_{t+1}$ follows an AR(1) process, $log\varepsilon_{t+1} = b \cdot log\varepsilon_t + u_{t+1}$, we get $C_{t+1} = S_{t+1} \cdot \prod_{i=0}^{\infty} exp(b^i \cdot u_{t+1-i})$ and, hence,

$$\Delta c_{t+1} = \frac{a}{1-\lambda} - \xi_{t+1} - \sum_{i=1}^{\infty} b^{i-1} \left(b - \left(1-\lambda\right)\right) \xi_{t+1-i}, \; 0 < \lambda < 1, \tag{29}$$

where $\xi_{t+1} \equiv -\frac{u_{t+1}}{1-\lambda}$.

As we saw above, when $\varphi = \gamma$ (the standard power utility model), $\sigma = \frac{1}{\gamma}$ for any values of λ and b. When $\varphi \neq \gamma$ and $b = 0$, the elasticity of intertemporal substitution is the reciprocal of the arithmetic average of the RRA coefficient γ and the parameter φ for any value of λ ($0 \leqslant \lambda \leqslant 1$):

$$\sigma = \frac{1}{(1-\lambda)\gamma + \lambda\varphi}. \tag{30}$$

When $\lambda = 0$, $S_{t+1} = exp\,(a)\, S_t$ and, hence, the SDF for model (10),

$$M_{t+1} = \delta^* \left(\frac{C_{k,t+1}}{C_{k,t}}\right)^{-\gamma}, \; \delta^* \equiv \delta \cdot exp\left(a\left(\gamma - \varphi\right)\right), \tag{31}$$

is observationally equivalent to that for the power utility model. So, it is not astonishing that with $\lambda = 0$, both models yield the same elasticity of intertemporal substitution, $\sigma = \frac{1}{\gamma}$. In another extreme case, when there is full adjustment in one period ($\lambda = 1$), $\sigma = \frac{1}{\varphi}$. It follows that under the assumption that the reference consumption level fully adjusts in one period, utility function (10) allows us not only to directly estimate the parameter of elasticity of intertemporal substitution, σ, but also completely disentangle risk aversion and elasticity of substitution.[12]

4 Empirical Analysis

Empirical evidence in Semenov (2002) is that the approximate equilibrium multifactor model for expected asset returns is able to explain both the equity premium and the return on the risk-free asset with economically plausible values of the RRA coefficient and

[12] There are two different parameters γ and φ which govern risk aversion and the elasticity of intertemporal substitution, respectively.

the time discount factor when all individuals are assumed to have the CRRA homogeneous preferences and the agent's marginal utility of consumption is expanded as a Taylor series up to cubic terms.

An undesirable property of the CRRA utility specification is that the elasticity of intertemporal substitution is constrained to be the reciprocal of the RRA coefficient. An attractive feature of the expected utility model with a reference level of consumption is that it allows to disentangle risk aversion and intertemporal substitution. GRS (2003a) test this utility function under the assumption of market completeness and find that this specification of preferences allows to obtain the point estimate of the elasticity of intertemporal substitution that is in the conventional range and statistically different from the inverse of the RRA coefficient.

In this section, we assume incomplete consumption insurance and limited participation and test the expected utility function with a reference level of consumption using the approximate equilibrium multifactor model for expected asset returns.

4.1 Description of the Data

The Consumption Data. The consumption data are taken from the CEX. As opposed to the PSID which offers only food consumption data on an annual basis, the CEX contains highly detailed data on monthly consumption expenditures.[13] The CEX attempts to account for an estimated 70 percent of total household consumption expenditures. Since the CEX is designed with the purpose of collecting consumption data, measurement error in consumption is likely to be smaller for CEX consumption data compared to the PSID consumption data.

The CEX data available cover the period from 1979:10 to 1996:2. It is a collection of data on approximately 5000 households per quarter in the United States. Each household in the sample is interviewed every three months over five consecutive quarters.[14] As households complete their participation, they are dropped and new households move into the sample. Thus, each quarter about 20 percent of the consumer units are new. The second through fifth interviews use uniform questionnaires to collect demographic and family characteristics as well as data on monthly consumption expenditures for the previous three months made by households in the survey.[15] Various income information is collected in the second and fifth interviews as well as information on the employment of each household member.

The measure of consumption used in this empirical investigation is consumption of

[13]Food consumption is likely to be one of the most stable consumption components. Furthermore, as Carroll (1994) points out, 95% of measured in the PSID food consumption is noise due to the absence of interview training.

[14]The first interview is practice and is not included in the published data set.

[15]Demographic variables are based upon heads of households.

nondurables and services (NDS). For each household, we calculate monthly consumption expenditures for all the disaggregate consumption categories offered by the CEX. Then, we deflate obtained values in 1982-84 dollars with the CPI's (not seasonally adjusted, urban consumers) for appropriate consumption categories.[16] Aggregating the household's monthly consumption across these categories is made according to the National Income and Product Account definitions of consumption aggregates. In order to transform my consumption data to a per capita basis, we normalize the consumption of each household by dividing it by the number of family members in the household.

The Returns Data. The measures of the nominal market return are the value-weighted and equal-weighted returns (capital gain plus dividends) on all stocks listed on the NYSE and AMEX obtained from the Center for Research in Security Prices (CRSP) of the University of Chicago. The real, monthly market return is calculated as the nominal market return divided by the 1-month inflation rate based on the deflator defined for NDS consumption. The nominal, monthly risk-free rate of interest is the 1-month Treasury bill return from CRSP. The real, monthly risk-free interest rate is calculated as the nominal risk-free rate divided by the 1-month inflation rate. The market premium is calculated as the difference between the real market return and the real risk-free rate of interest.

Asset Holders. For the consumer units completing their participation in the first through third quarters of 1986, BLS has changed, beginning the first quarter of 1986, the consumer unit identification numbers so that the identification numbers for the same household in 1985 (when this household has been interviewed for the first time) and in 1986 (when he has completed his interviews) are not the same. To match the consumer units between the 1985 and 1986 data tapes, we use the household characteristics which allow to identify consumer units uniquely. As a result, we manage to match 47.0% of households between the 1985 and 1986 data tapes. The detailed description of the procedure used to match the consumer units between the 1985 and 1986 data tapes is given in Semenov (2002).

In the fifth (final) interview, the household is asked to report end-of-period estimated market value of all stocks, bonds, mutual funds, and other such securities (market value of all securities) held by the consumer unit on the last day of the previous month as well as the difference in the estimated market value of all securities compared with the value of all securities held a year ago last month. Using these two values, we calculate asset holdings at the beginning of a 12-month recall period. The consumer unit is considered as an assetholder if the household's asset holdings at the beginning of a 12-month recall period

[16] The CPI's series are obtained from the BLS through CITIBASE.

exceed a certain threshold. To assess the quantitative importance of limited participation of households in the asset markets, we consider four sets of households. The first set (SET1) consists of all consumer units independently of the reported market value of all securities. To take into consideration that only a part of households participates in the asset markets, we use three sets of households defined as assetholders. The first one (SET2) consists of the consumer units whose asset holdings are equal to or exceed \$2 in 1999 dollars, the two others consist of the households reported total assets equal to or exceeding \$10000 (SET3) and \$20000 (SET4).[17]

Per capita consumption of a set of households is calculated as the equal-weighted average of normalized consumption expenditures of the households in the set. Obtained per capita consumption is seasonally adjusted by using the X-11 seasonal adjustment program.[18] We seasonally adjust the normalized consumption of each household by using the additive adjustments obtained from per capita consumption.

Data Selection Criteria. Following Vissing-Jorgensen (1998), we drop from the sample the bottom and the top percent of consumption growth observations for each month (under the assumption that these extreme values reflect reporting or coding errors). In addition, we drop nonurban households, households residing in student housing, households with incomplete income responses, and households who do not have a fifth interview. Following Brav, Constantinides, and Géczy (2002), in any given month, we drop from the sample households that report in that month as zero either their food consumption or their consumption of nondurables and services, or their total consumption, as well as households with missing information on the above items. Additionally, we keep in the sample only the households whose head is between 19 and 75 years of age.

[17] Over the period 1991-1996 about 18% of households, for which the market value of all securities held a year ago last month is not missing, reported asset holdings of \$1 at the beginning of a 12-month recall period. That occurs when the household reported owning securities without precising their value (see Vissing-Jorgensen (1998)). Following Vissing-Jorgensen (1998), we classify these households as nonassetholders.

[18] Ferson and Harvey (1992) point out that since the X-11 program uses of past and future information in the time-averaging it performs, this type of seasonal adjustment may induce spurious correlation between the error terms of a model and lagged values of the variables and, hence, may cause improper rejections of the model based on tests of overidentifying restrictions. As alternatives to using X-11 program, Brav and Géczy (1995) propose to use a simpler linear filter (Davidson and MacKinnon (1993)) or the Ferson-Harvey (1992) method of incorporating forms of seasonal habit persistence directly in the Euler equation.

4.2 The Estimation Methodology

When all investors have homogeneous preferences of the form (10), the Euler equations for the equity premium (8) and the risk-free rate (9) can be written as

$$E_t\left[\left(1+\sum_{n=1}^{N}\frac{1}{n!}(-1)^n\left(\prod_{l=0}^{n-1}(\gamma+l)\right)\frac{Z_{n,t+1}}{h_{t+1}^n}\right)\left(\frac{S_{t+1}}{S_t}\right)^{\gamma-\varphi}RP_{t+1}\right]=0 \tag{32}$$

and

$$\delta E_t\left[\left(1+\sum_{n=1}^{N}\frac{1}{n!}(-1)^n\left(\prod_{l=0}^{n-1}(\gamma+l)\right)\frac{Z_{n,t+1}}{h_{t+1}^n}\right)\left(\frac{S_{t+1}}{S_t}\right)^{\gamma-\varphi}R_{f,t+1}\right]=$$
$$\left(\frac{h_{t+1}}{h_t}\right)^{\gamma}\left(1+\sum_{n=1}^{N}\frac{1}{n!}(-1)^n\left(\prod_{l=0}^{n-1}(\gamma+l)\right)\frac{Z_{n,t}}{h_t^n}\right), \tag{33}$$

respectively.

Assuming the dynamics of the log reference level to be given by equation (16), we use the following two-stage procedure to estimate the parameters of interest. The first step is maximum likelihood (ML) estimation of the following regression model:

$$\Delta c_{t+1}=g-\eta_{t+1}+\theta\eta_t. \tag{34}$$

Given the estimates of g and θ obtained from (34), we are able to estimate the parameters a and λ of the behavioral model (22). The coefficient of η_t yields an estimate of $(1-\lambda)$ and, hence, of λ. The constant term g, when multiplied by $(1-\lambda)$, yields an estimate of a.

The second step is to use the iterated GMM approach to estimate the Euler equations for the premium of the real value-weighted and equal-weighted market portfolio returns over the risk-free rate as

$$E_t\left[\left(1+\sum_{n=1}^{N}\frac{1}{n!}(-1)^n\left(\prod_{l=0}^{n-1}(\gamma+l)\right)\frac{Z_{n,t+1}}{h^n}\right)\left(\prod_{i=0}^{I}\left(\frac{C_{t+1-i}}{C_{t-i}}\right)^{\lambda(1-\lambda)^i}\right)^{\gamma-\varphi}RP_{t+1}\right]=0 \tag{35}$$

and for the real risk-free rate as

$$\delta E_t\left[\left(1+\sum_{n=1}^{N}\frac{1}{n!}(-1)^n\left(\prod_{l=0}^{n-1}(\gamma+l)\right)\frac{Z_{n,t+1}}{h^n}\right)\left(\prod_{i=0}^{I}\left(\frac{C_{t+1-i}}{C_{t-i}}\right)^{\lambda(1-\lambda)^i}\right)^{\gamma-\varphi}R_{f,t+1}\right]=$$
$$1+\sum_{n=1}^{N}\frac{1}{n!}(-1)^n\left(\prod_{l=0}^{n-1}(\gamma+l)\right)\frac{Z_{n,t}}{h^n} \tag{36}$$

with λ replaced by its estimate obtained from (34).[19]

The sample period is from 1979:10 to 1996:2. As in Semenov (2002), we expand the agent's marginal utility of consumption as a Taylor series up to cubic terms ($N = 3$). The Euler equations for the excess value-weighted and equal-weighted market returns (35) and the Euler equation for the real risk-free interest rate (36) are estimated jointly using an iterated GMM approach. We exploit two sets of instruments. The first instrument set (INSTR1) consists of a constant, the real value-weighted and equal-weighted market returns, the real risk-free rate, and the real consumption growth rate lagged one period. The second set of instruments (INSTR2) is the first set extended with the same variables lagged an additional period.

4.3 Estimation Results

As in Semenov (2002), we assume the conditional expectation of consumption to be equal to the conditional expectation of aggregate consumption per capita, $h_{t+1} = E_t [C_{t+1}]$, and estimate the following random walk model of consumption:

$$\Delta c_{t+1} = g + \eta_{t+1}, \tag{37}$$

where $\Delta c_{t+1} \equiv log \frac{C_{t+1}}{C_t}$ and $\eta_{t+1} \sim N\left(0, \sigma_\eta^2\right)$. It follows that $h_{t+1} = exp\left(g + \frac{\sigma_\eta^2}{2}\right) C_t$. Table I presents the usual ML estimates for model (37).

The results of the ML estimation of (34) are reported in Table II. Neither the null hypothesis $H_0 : g = 0$ nor $H_0 : a = 0$ is rejected at the 5% level. The point estimates of λ are significantly different from both 0 and 1 at the 5% level for all the sets of households. The greatest value of λ is obtained for SET1, what means that the reference consumption level of non-assetholders adapts to changes in aggregate consumption per capita quicker than that of assetholders.

Given that the weight of consumption lagged ten periods becomes so small that one can neglect further past values of consumption, we estimate the Euler equations (35) and (36) with $I = 10$. The results of estimation and testing the Euler equations (35) and (36) with $N = 3$ are reported in Table III. Using the set of instruments INSTR1, we obtain

[19] It can be seen that when Abel's (1999) specification of consumption externalities is used, the parameters δ, G, and α_2 are not identifiable from (33). All we are able to identify is the parameter $\delta^* \equiv \delta G^{\alpha_2(\gamma-\varphi)}$. This leads to another problem. Given that $G \geqslant 1$ and $0 \leqslant \alpha_2 \leqslant 1$, the parameter δ cannot be estimated consistently when $G \neq 1$ and $\alpha_2 (\gamma - \varphi) \neq 0$. The estimate of δ is upward biased by the factor of $G^{\alpha_2(\gamma-\varphi)}$ when $G \neq 1$, $\alpha_2 \neq 0$, and $\gamma - \varphi > 0$ and downward biased when $G \neq 1$, $\alpha_2 \neq 0$, and $\gamma - \varphi < 0$. An attractive feature of our specification of the benchmark consumption level is that when $\lambda > 0$, the term $exp\left(a\left(\gamma - \varphi\right)\right)$ vanishes from the Euler equations as the ratio of benchmark levels in two successive periods is taken and, therefore, unlike Abel's (1999) specification of the reference consumption level, an unbiased estimate of δ can be obtained. Moreover, unlike Abel's (1999) specification, the presented in this paper specification of consumption externalities allows to estimate the growth rate of the benchmark level reflecting the increase in the standard of living (the parameter a in equation (34)).

the estimates of the RRA coefficient which are in the conventional range and significantly different from 0 for all the sets of households.[20] The point estimate of the elasticity of intertemporal substitution, σ, is positive only when a part of consumers is assumed to participate in the asset markets. However, only for SET2, σ is significantly positive at the 5% level. The standard power utility model $\left(H_0 : \gamma - \varphi = 0 \; (H_0 : \sigma = \frac{1}{\gamma})\right)$ is rejected for the households whose asset holdings are less than \$10000. The ratio model ($H_0 : \varphi = 1$) is rejected at the 5% significance level for all the households reported market value of all securities of less than \$20000. According to Hansen's J statistic, the model is not rejected statistically.

As the instrument set INSTR2 is used, we obtain the point estimates of γ which are significantly positive for all the sets. A little evidence of predictable variation in consumption growth in the face of predictable asset returns suggests that the elasticity of intertemporal substitution, σ, is small.[21] The results in Table III show that the point estimate of σ is small and significantly positive when only the households reported total assets equal to or exceeding \$10000 are classified as assetholders. Both the standard power utility specification and the ratio model are rejected statistically at the 5% significance level for all the sets of consumers. For SET1 and SET2, the point estimate of $\gamma - \varphi$ is significantly positive, what suggests that for the households reported total assets less than \$10000, consumption externalities are positive, while they are negative for the consumer units whose asset holdings are equal to or exceed \$10000 (for SET3 and SET4, the point estimate of $\gamma - \varphi$ is significantly negative). According to Hansen's test of the overidentifying restrictions, the model is not rejected statistically at the 5% level.

5 Concluding Remarks

The empirical results provide some evidence that the reference consumption level responses only gradually to changes in contemporaneous aggregate consumption per capita. The null hypotheses that the reference level only grows with the passage of time and that there is full adjustment in one period are both rejected statistically at the 5% significance level. This result is robust to the threshold value in the definition of assetholders. The rejection of the null hypothesis $\lambda = 1$ allows to conclude that Gali's specification of consumption externalities is not supported by the data. Empirical evidence is also presented that for nonassetholders, the reference level adapts to changes in aggregate consumption per capita more quickly than that for assetholders.

Another important result is that given partial adjustment of the reference level to

[20] The point estimates of γ are significantly different from 0 at the 5% significance level for SET1, SET2, and SET4. For SET3, the point estimate of risk aversion is significantly positive at the 10% level.

[21] See Campbell, Lo, and MacKinlay (1997) and Campbell and Mankiw (1990).

changes in aggregate consumption, we are able to disentangle risk aversion and intertemporal substitution. The obtained estimate of the elasticity of intertemporal substitution is in the conventional range and significantly different from the inverse of the RRA coefficient at the 5% level when the households reported the market value of all securities equal to or exceeding $10000 are classified as assetholders. Both the standard time-separable power utility model and the ratio preference specification are rejected statistically.

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Table I

Parameter Estimates for $\Delta c_{t+1} = g + \eta_{t+1}$

The sampling period is from 1979:10 to 1996:2. Four sets of households are considered. The first set (SET1) consists of all consumer units with any reported market value of all securities. We also use three sets of households classified as assetholders: SET2 consists of the households whose asset holdings are equal to or exceed $2 in 1999 dollars, the two others consist of the households reported total assets equal to or exceeding $10000 (SET3) and $20000 (SET4). The model is estimated by ML (standard errors in parentheses).

Parameters	SET1	SET2	SET3	SET4
g	0.0016	0.0022	0.0027	0.0021
	(0.0014)	(0.0023)	(0.0029)	(0.0032)
σ_η^2	0.0004	0.0010	0.0016	0.0020

Table II

Parameter Estimates for $\Delta c_{t+1} = g - \eta_{t+1} + \theta\eta_t$

The sampling period is from 1979:10 to 1996:2. Four sets of households are considered. The first set (SET1) consists of all consumer units with any reported market value of all securities. We also use three sets of households classified as assetholders: SET2 consists of the households whose asset holdings are equal to or exceed $2 in 1999 dollars, the two others consist of the households reported total assets equal to or exceeding $10000 (SET3) and $20000 (SET4). The model is estimated by ML (standard errors in parentheses). In Panel A, we report the values of the parameters estimated directly. The values of the parameters estimated indirectly are presented in Panel B. The standard errors for the parameters estimated indirectly are calculated by using the delta method.

Parameters	SET1	SET2	SET3	SET4
Panel A:				
g	0.0015	0.0020	0.0020	0.0018
	(0.0011)	(0.0014)	(0.0018)	(0.0021)
$\theta = 1 - \lambda$	0.2050	0.3486	0.3347	0.3401
	(0.0741)	(0.0736)	(0.0714)	(0.0698)
Panel B:				
a	0.0003	0.0007	0.0007	0.0006
	(0.0003)	(0.0005)	(0.0006)	(0.0007)
λ	0.7950	0.6514	0.6653	0.6599
	(0.0741)	(0.0736)	(0.0714)	(0.0698)

Table III

Parameter Estimates and Test Statistics for the Utility Specification with a Reference Level under Incomplete Consumption Insurance

The proxy for the reference consumption level growth rate is constructed using 10 lags of consumption growth. This has the effect of reducing the length of the sample by 10 months, so that the sampling period used in the estimation is from 1980:8 to 1996:2 rather than from 1979:10 to 1996:2. Four sets of households are considered. The first set (SET1) consists of all consumer units with any reported market value of all securities. We also use three sets of households classified as assetholders: SET2 consists of the households whose asset holdings are equal to or exceed \$2 in 1999 dollars, the two others consist of the households reported total assets equal to or exceeding \$10000 (SET3) and \$20000 (SET4). The agent's marginal utility of consumption is expanded as a Taylor series up to cubic terms ($N = 3$). The Euler equations for the excess value-weighted and equal-weighted market returns and for the real risk-free interest rate are estimated jointly using an iterated GMM approach (standard errors in parentheses). Two sets of instruments are exploited. The first instrument set (INSTR1) consists of a constant, the real value-weighted and equal-weighted market returns, the real risk-free rate, and the real consumption growth rate lagged one period. The second set of instruments (INSTR2) is the first set extended with the same variables lagged an additional period. The J statistic is Hansen's test of the overidentifying restrictions. The P value is the marginal significance level associated with the J statistic. In Panel A, we report the values of the parameters estimated directly from the Euler equations. The values of the parameters estimated indirectly are presented in Panel B. The standard errors for the parameters estimated indirectly are calculated by using the delta method.

Parameters	SET1	SET2	SET3	SET4
		INSTR1		
Panel A:				
γ	1.0912	1.2337	0.0238	1.1598
	(0.0201)	(0.0575)	(0.0125)	(0.0504)
φ	-46.6953	17.3727	0.2992	9.3589
	(3.0307)	(4.4998)	(0.2925)	(6.0644)
δ	0.9374	0.8701	0.9974	0.8976
	(0.0853)	(0.0509)	(0.0008)	(0.0954)
J statistic	6.3931	7.3585	8.0852	6.9422
P value	0.8950	0.8330	0.7784	0.8614
Panel B:				
σ	-0.0271	0.0851	4.8301	0.1522
	(0.0031)	(0.0230)	(4.5033)	(0.0937)
$\varphi - 1$	-47.6953	16.3727	-0.7008	8.3589
	(3.0307)	(4.4998)	(0.2925)	(6.0644)
$\gamma - \varphi$	47.7866	-16.1390	-0.2755	-8.1991
	(3.0313)	(4.4766)	(0.3003)	(6.0600)

Table III (continued)

Parameters	SET1	SET2	SET3	SET4
		INSTR2		
Panel A:				
γ	1.0887	1.0224	0.2528	0.3829
	(0.0130)	(0.0239)	(0.0538)	(0.0617)
φ	-33.6584	-21.2708	7.2052	12.5850
	(2.2501)	(2.6538)	(1.5513)	(2.1250)
δ	1.0737	0.9627	1.0218	0.9975
	(0.0324)	(0.0174)	(0.0054)	(0.0139)
J statistic	8.0897	8.7905	8.7992	8.6911
P value	0.9990	0.9980	0.9980	0.9982
Panel B:				
σ	-0.0377	-0.0741	0.2050	0.1186
	(0.0045)	(0.0131)	(0.0482)	(0.0231)
$\varphi - 1$	-34.6584	-22.2708	6.2052	11.5850
	(2.2501)	(2.6538)	(1.5513)	(2.1250)
$\gamma - \varphi$	34.7471	22.2932	-6.9524	-12.2021
	(2.2489)	(2.6565)	(1.5447)	(2.1108)