Risk premia and term premia in general equilibrium

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Received 11 August 1997; received in revised form 5 January 1998; accepted 7 January 1998

Abstract

The equity premium consists of a term premium reflecting the longer maturity of equity relative to short-term bills, and a risk premium reflecting the stochastic nature of equity payoffs and the deterministic nature of payoffs on riskless bills. This paper analyzes term premia and risk premia in a general equilibrium model with catching up with the Joneses preferences and a novel formulation of leverage. Closed-form solutions for moments of asset returns are derived. First-order approximations illustrate the effects of parameters and provide an algorithm to match the means and variances of the riskless rate and the rate of return on equity. © 1999 Elsevier Science B.V. All rights reserved.

JEL classification: G12

Keywords: Asset pricing; Equity premium; Risk premium; Term premium

1. Introduction

The average annual rate of return on equity exceeds the average annual rate of return on short-term riskless bills by several hundred basis points. There are two components to this equity premium: a risk premium and a term premium. The risk premium reflects the fact that equity is claim on stochastic payoffs, whereas a short-term riskless bill is a fixed-income security that is a claim on a known payoff. The term premium reflects the longer maturity of equity relative to short-term bills.

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The literature on the equity premium has typically focused on the overall equity premium rather than on the separate components. In the most influential quantitative application of the Lucas (1978) fruit-tree model of asset pricing, Mehra and Prescott (1985) found that this model was incapable of producing an equity premium of more than 35 basis points when they confined attention to parameter values that they deemed plausible. With such a small overall equity premium implied by the model, it did not much matter how much was a risk premium and how much was a term premium. However, subsequent research has produced models that can account for an equity premium of several hundred basis points per year. I will show in this paper how the equity premium in a general equilibrium model can be decomposed into a risk premium and a term premium. This decomposition provides helpful insights about the source of the equity premium and provides guidance in choosing appropriate parameter values.

The most basic application of the Lucas fruit-tree model can be used to price unlevered equity in an economy with i.i.d. consumption growth and a representative consumer with time-separable isoelastic utility. The literature has extended this basic model in many ways, and I will adopt two of these extensions, though I will combine them in a way to obtain fresh insights. First, I will specify preferences to have the catching up with the Joneses feature introduced in Abel (1990) and used more recently by Campbell and Cochrane (1994) and Carroll et al. (1997). Second, I will introduce a novel tractable formulation of leverage. Despite the additional richness introduced by these extensions, I derive closed-form solutions for expected rates of return, term premia, and risk premia for a general class of assets.

The catching up with the Joneses feature of the utility function was originally introduced to help account for the high average value of the equity premium observed empirically. However, when this form of the utility function was specified to imply a realistic value of the equity premium in Abel (1990), the model produced a riskless rate of return that was far too volatile. Campbell and Cochrane (1994) developed a form of catching up with the Joneses preferences that yielded, as in actual data, a large equity premium and low variability of the riskless rate. They achieved this low variability of the riskless rate by specifying a complicated recursive function for the determination of the benchmark level of consumption. Here I adopt a simpler formulation of the benchmark level of consumption that produces, with the inclusion of leverage, low variability of the riskless rate along with a large equity premium.

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1 Recently, Jermann (1998) has examined term premia and payout uncertainty premia in a model with preferences that exhibit habit formation rather than catching up with the Joneses as in the current paper. Boldrin et al. (1995) also examine asset pricing with habit formation, and Lettau and Uhlig (1995) examine asset pricing behavior with a form of catching up with the Joneses utility, but unlike Jermann (1996) neither of these papers examines term premia.
The analysis of leverage arises naturally from the formulation of the canonical asset introduced in this paper. The payoff in period $t$ on the canonical asset is specified to be proportional to $y_t^\lambda$, where $y_t$ is an observable random variable and $\lambda$ is a constant. I use this formulation in order to include fixed-income securities and equities as special cases. For fixed-income securities, $\lambda = 0$ so that payoffs are nonstochastic. Unlevered equity in the Lucas fruit-tree model is modeled by setting $\lambda = 1$ and $y_t$ equal to aggregate output (which equals consumption) per capita. Although the form of the canonical asset was initially developed with only the values of $\lambda = 0$ and $\lambda = 1$ in mind, it became apparent that much of the analysis can be conducted, and closed-form solutions derived, for arbitrary values of $\lambda$. Indeed, values of $\lambda$ greater than one provide a good approximation to levered equity.

Although closed-form solutions are derived for various moments of asset returns, many of these expressions are too cumbersome to clearly illustrate the effects of various parameters on the moments of returns. To make these effects transparent, I derive first-order approximations to the exact solutions. Numerical calibration demonstrates that these approximations yield values close to the values calculated using the exact solutions. More importantly, I use these approximations to develop a simple algorithm for choosing parameter values that allow the approximate unconditional means and variances of the riskless rate and the rate of return on equity to match the corresponding historical sample values.

Section 2 develops the model with catching up with the Joneses preferences and introduces the canonical asset. Asset pricing is discussed in Section 3 which includes a discussion of risk premia and term premia without restricting the distribution of growth rates. Beginning in Section 4, I assume that growth rates are lognormal. The interpretation of the parameter $\lambda$ as a measure of leverage is discussed in Section 5. In Section 6, I derive closed-form expressions for the means and variance of the riskless rate and the rate of return on equity. I calibrate the model in Section 7 and develop the algorithm for choosing parameter values that allow the model’s predictions of the unconditional means and variances of the riskless rate and the rate of return on levered equity to match the corresponding empirical moments. Concluding remarks are presented in Section 8.

2. The model

Consider a closed economy populated by a continuum of identical infinitely-lived consumers. Output in this economy is a homogeneous good that is completely perishable. In equilibrium, all output is consumed in the period in which it is produced so that, as in the Lucas (1978) fruit-tree model,
consumption equals output. The amount of output per person in period $t$ is $C_t > 0$, which equals the consumption of the representative consumer.

2.1. Preferences

In period $t$ an individual consumer chooses a level of consumption $c_t$ to maximize utility $U_t$, which is given by the function

$$U_t = E_t \left\{ \sum_{j=0}^{\infty} \left( \frac{1}{1+\delta} \right)^j u(c_{t+j}, v_{t+j}) \right\},$$

where $u(c_t, v_t) = \frac{1}{1-\alpha} \left( \frac{c_t}{v_t} \right)^{1-\alpha}$, $\alpha > 0$, $\delta > 0$ (1)

and $v_t$ is a benchmark level of consumption which is exogenous to an individual consumer.\(^2\) The curvature parameter $\alpha$ is the coefficient of relative risk aversion. If $v_t$ is a fixed constant, then the utility function is the standard time-separable utility function with a constant coefficient of relative risk aversion $\alpha$ and a constant rate of time preference $\delta$.

The benchmark level of consumption, $v_t$, tends to grow over time as the standard of living in a country rises. Specifically,

$$v_t = C_t^\gamma_0 C_{t-1}^\gamma_1 (G_t)^{\gamma_2},$$

(2)

where $G \geq 1$ and $0 \leq \gamma_i \leq 1$ for $i = 0, 1, 2$. The dependence of the benchmark level of consumption on current and lagged aggregate consumption per capita, $C_t$ and $C_{t-1}$, is a generalization of the ‘catching up with the Joneses’ specification of utility introduced in Abel (1990), and captures the notion that the benchmark level of consumption is an increasing (homogeneous) function of current and recent levels of consumption per capita. The special case with $\gamma_1 > 0$ and $\gamma_0 = \gamma_2 = 0$ corresponds the simple formulation of catching up with the Joneses in Abel (1990). The case with $\gamma_0 > 0$ and $\gamma_1 = \gamma_2 = 0$ corresponds to Gali’s (1994) specification of consumption externalities. The dependence of the benchmark level of consumption on $G_t$ allows for the possibility that the benchmark level of consumption grows simply with the passage of time.

\(^2\) Campbell and Cochrane (1994) assume that the representative consumer’s utility depends on the difference between $c_t$ and $v_t$ rather than on the ratio of $c_t$ to $v_t$. Specifically, they assume that $u(c_t, v_t) = (c_t - v_t)^{1-\alpha}/(1-\alpha)$. With the formulation in Campbell and Cochrane, care must be taken to prevent $c_t$ from falling below $v_t$. The formulation in Eq. (1) makes this consideration unnecessary.
Equilibrium asset prices and returns depend on the marginal rate of substitution of the representative consumer. For the utility function in Eq. (1) and the benchmark level of consumption in Eq. (2) the marginal rate of substitution between period \( t \) and period \( t + 1 \) is

\[
M_{t+1} = \frac{1}{1 + \delta} \frac{u(c_{t+1}, v_{t+1})}{u(c_t, v_t)}
\]

\[
= \frac{1}{1 + \delta} \left( \frac{c_{t+1}}{c_t} \right)^{-\gamma} \left( \frac{C_{t+1}}{C_t} \right)^{\gamma_0} \left( \frac{C_t}{C_{t-1}} \right)^{\gamma_1} \left( G^{\gamma_2} \right)^{\gamma_2 - 1}.
\]

(3)

The marginal rate of substitution can be expressed more conveniently in terms of the growth rate of consumption. Let \( x_{t+1} \) be the ratio of consumption in period \( t + 1 \) to consumption in period \( t \) and observe that in equilibrium

\[
x_{t+1} \equiv \frac{C_{t+1}}{C_t} \equiv \frac{c_{t+1}}{c_t}.
\]

(4)

Substituting Eq. (4) into Eq. (3) gives an expression for the equilibrium marginal rate of substitution

\[
M_{t+1} = \beta x_{t+1}^A x_t^\theta
\]

(5a)

where

\[
\beta \equiv \frac{G^{\gamma_2 (x-1)}}{1 + \delta} > 0,
\]

(5b)

\[
A \equiv x - \gamma_0 (x - 1) > 0,
\]

(5c)

and

\[
\theta \equiv \gamma_1 (x - 1).
\]

(5d)

Although the specification of preferences involves six parameters – \( \delta, x, \gamma_0, \gamma_1, \gamma_2 \) and \( G \) – and all six of the parameters affect the marginal rate of substitution, there are only three independent parameters that determine the marginal rate of substitution. Thus, relative to the standard time-separable isoelastic specification of preferences, which has two independent parameters – the rate of time preference and the coefficient of relative risk aversion – the current specification of preferences introduces only one additional degree of freedom. The fact that there are six parameters but only three independent parameters allows for a variety of interpretations of preferences. However, from the perspective of
fitting empirical data on asset returns, there are only three preference parameters: \( \beta, A \) and \( \theta \).

### 2.2. The canonical asset

Rather than proceed with separate derivations for the prices and rates of return for different assets such as equity, short-term bills and long-term bonds, I will introduce a canonical asset that includes all of these assets as special cases. The canonical asset is an \( n \)-period asset. From the standpoint of period \( t \), the terminal period of an \( n \)-period asset is period \( t + n \). This asset pays \( a_j y_{t+n-j} \) in the period that is \( j \) periods before the terminal period for \( j = 0, \ldots, n - 1 \), where \( y_{t+n-j} > 0 \) is a random variable, \( a_0 > 0 \) is a constant, \( a_j \geq 0, j = 1, \ldots, n - 1 \) are constants, and \( \lambda \) is a constant that indexes the variability of future payoffs. Thus, for instance, in the terminal period, the asset pays \( a_0 y_{t+n} \) and in period \( t + 1 \) the asset pays \( a_{n-1} y_{t+1}^2 \).

The canonical asset introduced here includes equities and fixed-income securities of all maturities. Fixed-income securities such as bonds and bills are represented by \( \lambda = 0 \) which implies that the payoff in period \( t + n - j \) is the known amount \( a_j \). A standard coupon bond with face value \( F \) and coupon \( d \) is represented by \( a_0 = F + d \), and \( a_1 = \cdots = a_{n-1} = d > 0 \); in this formulation, a pure discount bond is represented by \( d = 0 \).

Securities with risky payoffs have nonzero values of \( \lambda \). For instance, in the Lucas (1978) fruit-tree model, the dividend (per capita) on unlevered equity equals consumption per capita \( C_t \). In terms of our canonical asset, unlevered equity in the Lucas model is an infinite-period asset that pays \( C_t \) in period \( t \). Using the notation for canonical assets, \( n = \infty \), \( a_j = 1 \) for all \( j \geq 0 \), \( y_t \equiv C_t \), and \( \lambda = 1 \). As discussed in Section 5, levered equity can be represented by values of \( \lambda \) greater than one.

Let \( p(n, \lambda) \) be the ex-payment price of the canonical \( n \)-period asset in period \( t \). The price of this asset also depends on the sequence of constants \( a_j, j = 0, \ldots, n - 1 \), and on the stochastic specification of \( y_t \), but this dependence

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3These three parameters can be used to state the condition for utility in Eq. (1) to be finite. Specifically, if \( x_t \) is i.i.d. over time, then utility in Eq. (1) is finite if and only if \( \beta E[x^{1-A+\theta}] < 1 \). To prove this result, rewrite Eq. (1) as \( U_t = \sum_{j=0}^{\infty} \tilde{u}_{t+j} \) where \( \tilde{u}_{t+j} = E_t \left\{ \left( \frac{1}{1 + \delta} \right)^n u_{t+j+n} \right\} \). In equilibrium,

\[
\tilde{u}_{t+j} = \frac{[G; \delta; x^{-1}]/(1 + \delta)]E_t[C_{t+j+1}^{-A}C_{t+j}^{\theta}]}{E_t[C_{t+j+1}^{-A}C_{t+j}^{\theta}]} = \frac{\beta E_t[x_{t+j}^{-A}C_{t+j+1}^{-A+\theta}]}{E_t[x_{t+j}^{-A}C_{t+j+1}^{-A+\theta}]}.
\]

If \( x_t \) is i.i.d. over time, then for \( j \geq 1 \), \( \tilde{u}_{t+j} = \beta E[x^{1-A+\theta}] \). Therefore, \( U_t = \sum_{j=0}^{\infty} \tilde{u}_{t+j} \) is finite if and only if \( \beta E[x^{1-A+\theta}] < 1 \).
is not reflected in the notation. The gross rate of return on the canonical asset between period \( t \) and period \( t+1 \) is
\[
R_{t+1}(n, \lambda) = \frac{p_{t+1}(n-1, \lambda) + a_{n-1}y_{t+1}^\lambda}{p_t(n, \lambda)}, \quad \text{for } n \geq 1. \tag{6}
\]

The moments of the rate of return on the canonical asset can be determined after specifying the stochastic structure of the economy.

2.3. The stochastic structure

Recall that \( x_{t+1} \equiv C_{t+1}/C_t \) is the growth rate of aggregate consumption per capita. Define \( z_{t+1} \equiv y_{t+1}/y_t \) to be the growth rate of \( y_t \). As noted earlier, in the specification of unlevered equity in the Lucas fruit-tree model, \( y_t \equiv C_t \) so that \( z_{t+1} \equiv x_{t+1} \). More generally, \( x_{t+1} \) and \( z_{t+1} \) are distinct random variables and are elements of a random vector of growth rates \( W_{t+1} \) which is observable at the beginning of period \( t+1 \). I assume that \( W_{t+1} \) is i.i.d.

3. Asset pricing

In this section I use the marginal rate of substitution in Eq. (5a) to determine the price of the canonical asset. As will become evident, the price of the canonical asset can be expressed as
\[
p_t(n, \lambda) = \omega(n, \lambda)x_t^\theta y_t^\lambda, \tag{7}
\]
where \( \omega(n, \lambda) \) is a function to be determined. Substituting Eq. (7) into Eq. (6) and recalling that \( z_{t+1} \equiv y_{t+1}/y_t \) yields
\[
R_{t+1}(n, \lambda) = \frac{\omega(n-1, \lambda)x_{t+1}^\theta + a_{n-1}y_{t+1}^\lambda}{\omega(n, \lambda)x_t^\theta}z_{t+1}^\lambda, \quad \text{for } n \geq 1. \tag{8}
\]

It is well known that the product of the gross rate of return and the intertemporal marginal rate of substitution has a conditional expectation equal to one. In the notation of this paper
\[
E_t\{R_{t+1}(n, \lambda)M_{t+1}\} = 1. \tag{9}
\]
Substituting Eqs. (5a) and (8) into Eq. (9), and using the assumption that \( W_{t+1} \) is i.i.d., yields the following difference equation for \( \omega(n, \lambda) \)
\[
\omega(n, \lambda) = \kappa(\lambda)\omega(n-1, \lambda) + \beta d_{n-1}E_t\{x_t^{-A}z_{t+1}^\lambda\}, \quad \text{for } n \geq 1. \tag{10a}
\]
where
\[
\kappa(\lambda) \equiv \beta E_t\{x_t^{-A}z_{t+1}^\lambda\}. \tag{10b}
\]
I will assume that $\beta$ and the distribution of the vector $W_{t+1}$ are such that $0 < \kappa(\lambda) < 1$. This assumption guarantees that the difference equation converges as $n$ grows.\textsuperscript{4}

The boundary condition for this difference equation is provided by the fact that after the asset has yielded its last payment, and is formally a zero-period asset, its price is zero. Thus the boundary condition is $p_t(0, \lambda) = 0$ which implies $\omega(0, \lambda) = 0$. Substituting this boundary condition into Eq. (10a) yields

$$\omega(1, \lambda) = \beta a_0 E\{x^{-A} \sigma^2 \} > 0.$$  \hfill (11)

It is straightforward to verify that the following expression for $\omega(n, \lambda)$ satisfies the difference Eq. (10a) and the boundary condition (11)

$$\omega(n, \lambda) = \frac{\omega(1, \lambda)}{a_0} \sum_{i=1}^{n} a_{i-1} [\kappa(\lambda)]^{n-i}.$$ \hfill (12)

The function $\omega(n, \lambda)$ can be either increasing or decreasing in the maturity $n$.

- For a pure discount security, which is characterized by $a_1 = \cdots = a_{n-1} = 0$, Eq. (12) implies $\omega(n, \lambda) = \omega(1, \lambda)[\kappa(\lambda)]^{n-1}$ which is strictly decreasing in $n$.

  Thus, for given values of $x_t$ and $y_t$, the price of an $n$-period discount security, $p_t(n, \lambda) = \omega(n, \lambda)x_t^\theta y_t^\lambda$ is strictly decreasing in maturity $n$.

- For a security with $a_0 = a_1 = \cdots = a_{n-1} > 0$, Eq. (12) implies $\omega(n, \lambda) = \omega(1, \lambda) \frac{1 - [\kappa(\lambda)]^n}{1 - \kappa(\lambda)}$, which is strictly increasing in $n$.

\textsuperscript{4}The assumption that $\kappa(\lambda) < 1$ also leads to a sufficient condition to guarantee that utility in Eq. (1) is finite in the case in which $z \equiv x$ and $\ln x$ is $N(\mu, \sigma^2)$, which is widely considered in the literature and is considered in Section 6 and Section 7. In this case, $\kappa(\lambda) \equiv \beta E\{x^{1-A+\theta} \}$ which implies

$$\beta E\{x^{1-A+\theta} \} = \beta E\{x^{1-A+\theta} \} \exp[-(\lambda - 1) \left( \mu + \left( \theta - A + \frac{1 + \lambda}{2} \right) \sigma^2 \right)]$$

Therefore, the assumption that $\kappa(\lambda) < 1$ implies that

$$\beta E\{x^{1-A+\theta} \} < 1 \text{ if } (\lambda - 1) \left( \mu + \left( \theta - A + \frac{1 + \lambda}{2} \right) \sigma^2 \right) \geq 0.$$  

Thus, the assumption that $\kappa(\lambda) < 1$ implies that

$$(\lambda - 1) \left( \mu + \left( \theta - A + \frac{1 + \lambda}{2} \right) \sigma^2 \right) \geq 0$$

is a sufficient condition for utility in Eq. (1) to be finite when $z \equiv x$. See footnote 3.
• For a coupon bond with face value $F$ and coupon $d$, $a_0 = F + d$ and $a_1 = \cdots = a_{n-1} = d > 0$. In this case, Eq. (11) implies

$$\omega(n, \lambda) = \frac{\omega(1, \lambda)}{F + d} \left[ (\kappa(\lambda))^n - 1 \right] F + \frac{1 - (\kappa(\lambda))^n}{1 - \kappa(\lambda)} d,$$

which is increasing in maturity $n$ if $d/F > (1 - \kappa(\lambda))/\kappa(\lambda)$ but is decreasing in $n$ if $d/F < (1 - \kappa(\lambda))/\kappa(\lambda)$. Thus, the price of an $n$-period coupon bond is an increasing, decreasing, or constant function of the maturity $n$ depending on whether $d/F$ is greater than, less than, or equal to $(1 - \kappa(\lambda))/\kappa(\lambda)$.

3.1. The expected rate of return on the canonical asset

Empirical and theoretical studies of asset pricing commonly focus on the expected rates of return on various assets such as equities and fixed-income securities. Before computing the expected rate of return on the general canonical asset, it is easiest to begin with the expected rate of return on one-period assets. Taking the expectation of both sides of Eq. (8), setting $n = 1$, and using Eq. (11) yields

$$E[R_{t+1}(1, \lambda)] = \frac{E[z_{t+1}^2]}{\beta E[x_{t+1}^2 + z_{t+1}^2] x_t^2}. \tag{13}$$

It can be shown that

$$E[R_{t+1}(n, \lambda)] = \left[ \Psi + (1 - \Psi) \frac{a_{n-1}}{a_0} \frac{\omega(1, \lambda)}{\omega(n, \lambda)} \right] E[R_{t+1}(1, \lambda)], \tag{14a}$$

5 Taking the expectation of both sides of Eq. (8) for maturity equal to $n$ and then for maturity equal to 1 yields

$$\frac{E[R_{t+1}(n, \lambda)]}{E[R_{t+1}(1, \lambda) = \frac{\omega(n - 1, \lambda) E[x_{t+1}^\lambda z_{t+1}^\lambda] + a_{n-1} E[z_{t+1}^\lambda]}{\omega(n, \lambda) a_0 E[z_{t+1}^\lambda}$$

which can be rewritten as

$$\frac{E[R_{t+1}(n, \lambda)]}{E[R_{t+1}(1, \lambda)]} = \frac{\omega(n - 1, \lambda) E[x_{t+1}^\lambda z_{t+1}^\lambda] \omega(1, \lambda)}{\omega(n, \lambda) a_0 E[z_{t+1}^\lambda} + a_{n-1} \omega(1, \lambda)}$$

Now use Eq. (10a) to obtain

$$\frac{E[R_{t+1}(n, \lambda)]}{E[R_{t+1}(1, \lambda)]} = \frac{1}{\kappa(\lambda)} \left[ 1 - \frac{\beta a_{n-1} E[x_{t+1}^\lambda z_{t+1}^\lambda]}{\omega(n, \lambda) a_0 E[z_{t+1}^\lambda} \right] \frac{E[x_{t+1}^\lambda z_{t+1}^\lambda] \omega(1, \lambda)}{a_0 \omega(n, \lambda)} + a_{n-1} \omega(1, \lambda).$$

Next use Eq. (11), the definition of $\kappa(\lambda)$ in Eq. (10b), and the definition of $\Psi$ in Eq. (14b) to obtain

$$\frac{E[R_{t+1}(n, \lambda)]}{E[R_{t+1}(1, \lambda)]} = \left[ 1 - \frac{a_{n-1} \omega(1, \lambda)}{a_0 \omega(n, \lambda)} \right] \Psi + \frac{a_{n-1} \omega(1, \lambda)}{a_0 \omega(n, \lambda)}$$

which is equivalent to Eq. (14a).
In a somewhat different model, Backus, Gregory, and Zin (1989) show (Proposition 1, p. 382) that term premia will be identically zero if the marginal rate of substitution is independent over time.

In the model presented here with \( x_t \) i.i.d. over time, if \( \theta = 0 \), the marginal rate of substitution, \( M_{t+1} = \beta x_{t+1}^{-\lambda} \), is independent over time; the term premium will be zero as in Backus, Gregory, and Zin.

3.2. Term premia

For any given value of \( \lambda \) and any given choice of the variable \( y_t \) (which determines the payoff \( a_j y_{t+n-j} \)), there is a term structure of expected one-period rates of return on assets of various maturities. The term premium is the excess of the expected one-period rate of return on a \( n \)-period asset over the expected one-period rate of return on a one-period asset with the same value of \( \lambda \). Formally, the term premium on an \( n \)-period asset is

\[
\text{TP}(n, \lambda) \equiv \frac{E[R_{t+1}(n, \lambda)]}{E[R_{t+1}(1, \lambda)]} - 1. \tag{15}
\]

Using the definition of the term premium in Eq. (15), it follows from Eq. (14a) that the term premium is

\[
\text{TP}(n, \lambda) = (\Psi - 1) \Gamma(n, \lambda), \tag{16a}
\]

where

\[
\Gamma(n, \lambda) \equiv 1 - \frac{a_{n-1}}{a_0} \left[ \frac{\omega(1, \lambda)}{\omega(n, \lambda)} \right] = 1 - \frac{a_{n-1}}{\sum_{i=1}^{n} a_{i-1}[\kappa(\lambda)]^{n-i}}. \tag{16b}
\]

According to Eq. (16a), the term premium is the product of two factors: \( \Psi - 1 \) and \( \Gamma(n, \lambda) \). The first factor, \( \Psi - 1 \), is independent of the maturity \( n \) and of the sequence of payoff coefficients \( a_j, j = 0, \ldots, n-1 \). Thus, I will call \( \Psi - 1 \) the ‘term premium scale factor’ because a change in this factor changes the term premia at all maturities by the same proportion. Inspection of the definition of \( \Psi \) in Eq. (14b) reveals that when \( \theta = 0 \), \( \Psi = 1 \) and hence the term premium scale factor is zero. Since \( \theta = \gamma_1(\alpha - 1) \), the term premium will be zero if utility is logarithmic (\( \alpha = 1 \)) or if \( \gamma_1 = 0 \). Thus, the term premium will be nonzero only if

6 In a somewhat different model, Backus, Gregory, and Zin (1989) show (Proposition 1, p. 382) that term premia will be identically zero if the marginal rate of substitution is independent over time. In the model presented here with \( x_t \) i.i.d. over time, if \( \theta = 0 \), the marginal rate of substitution, \( M_{t+1} = \beta x_{t+1}^{-\lambda} \), is independent over time; the term premium will be zero as in Backus, Gregory, and Zin.
the consumption externality, i.e., the dependence of an individual’s utility on aggregate consumption per capita, has a lagged component as reflected in $\gamma_1 > 0$. The contemporaneous component of the consumption externality parameterized by $\gamma_0$ does not affect whether the term premium is nonzero. Therefore, in the formulation of consumption externalities used by Gali (1994) with $\gamma_0 > 0$ and $\gamma_1 = 0$, there will be no term premium.

The second factor, $\Gamma(n, \lambda)$, captures the shape of the term structure for given $\lambda$. Notice that $\Gamma(1, \lambda) = 0$, and $\Gamma(n, \lambda) > 0$ for $n > 1$.7

- For an $n$-period security that makes no payments until the final period, such as an $n$-period discount bond, $a_1 = \cdots = a_{n-1} = 0$, and hence $\Gamma(n, \lambda) = 1$ for $n > 1$. Thus, the term premium is independent of the maturity $n$ for discount securities with a given $\lambda$ and with more than one period to run.
- For a security with $a_0 = a_1 = \cdots = a_{n-1} > 0$, $\Gamma(n, \lambda) = (\kappa(\lambda) - [\kappa(\lambda)]^n)/ (1 - [\kappa(\lambda)]^n)$ which is increasing in $n$.
- For a standard coupon bond with face value $F$ and coupon $d$, $a_0 = F + d$ and $a_1 = \cdots = a_{n-1} = d > 0$, so $\Gamma(n, \lambda) = 1 - \frac{d}{[\kappa(\lambda)]^{n-1}F + \frac{1 - [\kappa(\lambda)]^n}{1 - \kappa(\lambda)}d}$. The term premium is decreasing in maturity $n$ if $d/F < (1 - \kappa(\lambda))/\kappa(\lambda)$ and is increasing in maturity $n$ if $d/F > (1 - \kappa(\lambda))/\kappa(\lambda)$.

3.3. Risk premia

The risk premium is the expected excess rate of return on a risky asset relative to the rate of return on a riskless asset of the same maturity. I will define the risk premium in terms of one-period assets. A riskless one-period security is characterized by $n = 1$ and $\lambda = 0$, and has a known return $R_{t+1}(1, 0)$, so the risk premium is

$$
RP_t(1, \lambda) = \frac{E_t[R_{t+1}(1, \lambda)]}{R_{t+1}(1, 0)} - 1. 
$$

(17)

The rate of return on a one-period riskless security is easily calculated by setting $\lambda = 0$ in Eq. (13) to obtain

$$
R_{t+1}(1, 0) = \frac{1}{\beta E_t[x_{t+1}^{-\delta}]x_t^\delta}. 
$$

(18)

---

7Since $0 \leq a_{n-1} < \sum_{i=1}^{n-1}a_{i-1}[\kappa(\lambda)]^{n-i}$, Eq. (16b) implies that $0 < \Gamma(n, \lambda) \leq 1$. 

---
Now substitute Eqs. (13) and (18) into Eq. (17) to obtain

\[ RP_t(1, \lambda) = \frac{E_x \{ x_{t+1}^{-A} \} E_z \{ z_{t+1}^{1} \}}{E_x \{ x_{t+1}^{-A} z_{t+1}^{1} \}} - 1. \] (19)

This risk premium in Eq. (19) equals \(-\text{Cov} \{ x_{t+1}^{-A}, z_{t+1}^{1} \} / E_x \{ x_{t+1}^{-A} z_{t+1}^{1} \}\) so that the sign of the risk premium is the opposite of the sign of the conditional covariance of \( x_{t+1}^{-A} \) and \( z_{t+1}^{1} \). In the standard Lucas fruit-tree model where \( z_{t+1} = x_{t+1} \), this conditional covariance is negative and the risk premium is necessarily positive.\(^8\)

4. Lognormality

So far I have examined the expected rate of return on the canonical asset without imposing restrictions on the distribution of growth rates \( (W_{t+1}) \) except that \( \kappa(\lambda) \equiv \beta E_x \{ x_{t+1}^{-A} z_{t+1}^{1} \} < 1 \). Many contributions to the asset pricing literature assume that the distribution of growth rates is lognormal, and I will adopt this assumption for the remainder of the paper. Specifically, I assume that

\[
\begin{bmatrix}
\ln x_{t+1} \\
\ln z_{t+1}
\end{bmatrix}
\sim N\left(\begin{bmatrix}
\mu_x \\
\mu_z
\end{bmatrix}, \begin{bmatrix}
\sigma_x^2 & \sigma_{xz} \\
\sigma_{xz} & \sigma_z^2
\end{bmatrix}\right).
\] (20)

Under the distributional assumption in Eq. (20), the term premium scale factor \( \Psi \) defined in Eq. (14b) simplifies to

\[ \Psi = \exp(\theta A \sigma_x^2) \]. (21)

Although \( \Psi \) is defined in terms of the joint conditional distribution of \( z_{t+1} \) and \( x_{t+1} \), the value of \( \Psi \) depends only on the conditional variance of \( x_{t+1} \); it is independent of the parameters of the marginal conditional distribution of \( z_{t+1} \), and is also independent of the parameter \( \lambda \) which measures the variability of future payoffs. Thus, the term premium scale factor, \( \Psi - 1 \), depends on the preference parameters \( \theta \) and \( A \) and the variance of consumption growth \( \sigma_x^2 \), but it is independent of the payoff characteristics of the asset.

Substituting Eq. (21) into Eq. (16a) yields

\[ TP_t(n, \lambda) = \Gamma(n, \lambda)[\exp(\theta A \sigma_x^2) - 1] \cong \Gamma(n, \lambda)\theta A \sigma_x^2. \] (22)

---

\( ^8 \) This statement is based on the assumption that \( \lambda > 0 \). Unlevered equity in the fruit-tree model is represented by \( \lambda = 1 \), and, as we show in Section 5, levered equity is represented by \( \lambda > 1 \). Of course, if \( \lambda = 0 \), the asset is riskless and the risk premium in Eq. (19) is zero.
Eq. (22) presents an exact expression for the term premium as well as an approximation. According to the approximation, the term structure of term premia is proportional to $\theta A\sigma^2_x$, which is approximately equal to the term premium scale factor. Recall that $\Gamma(n, \lambda) > 0$ for $n > 1$ so that the term premium has the same sign as the term premium scale factor. The sign of the term premium scale factor is the same as the sign of $\theta$. Thus for $\theta > 0$, the term premium is positive for $n > 1$. However, if $\theta < 0$ (for instance, if $\gamma_1 > 0$ and $\tau < 1$), the term premium is negative for all $n > 1$.9

The assumption of conditional lognormality also leads to a simple expression for the risk premium. Under the distributional assumption in Eq. (20), the expression for the risk premium in Eq. (19) becomes

$$RP_t(1, \lambda) = \exp(\lambda \sigma_{xz}^2) - 1 \cong A\sigma_{xz}.$$

(23)

According to the approximation in Eq. (23) the risk premium equals $A$ multiplied by the covariance of $\ln x_{t+1}$ and $\lambda \ln z_{t+1}$ where $\ln x_{t+1}$ is the (logarithm of the) growth rate of consumption and $\lambda \ln z_{t+1}$ is the (logarithm of the) growth rate of the risky payoff on the asset. In the special case of the Lucas fruit-tree model, the payoff on unlevered equity is simply per capita consumption, so that $z_{t+1} \equiv x_{t+1}$. In this case, $\lambda = 1$ and $\sigma_{xz}^2 = \sigma^2_x$, so that the risk premium is approximately equal to $A\sigma^2_x$.

5. Leverage

Having shown in Eq. (23) that the (approximate) risk premium is proportional to $\lambda$, I will now show that the parameter $\lambda$ provides a convenient way to model leverage. Recall that if $\lambda = 0$, future payoffs are deterministic and are given by the sequence of constants $a_j, j = 0, 1, \ldots, n - 1$. Thus, with $\lambda = 0$, the canonical asset is a fixed-income security such as a bill or a bond. Alternatively, if $\lambda = 1$ and if $y_t \equiv C_t$ so that $z_{t+1} \equiv x_{t+1}$, the canonical asset is unlevered equity in a fruit-tree model. I will show below that levered equity can be represented approximately by values of $\lambda$ greater than one, with higher values of $\lambda$ representing higher degrees of leverage.

5.1. A one-period levered asset

Consider a one-period security in period $t$ that offers a stochastic payoff $\xi_{t+1}$ in period $t + 1$. Call this security $L$ (for ‘levered equity’) and let $PL_t$ denote

---

9 These findings are consistent with the result in Backus et al. (1989), (p. 382) that the sign of the term premium will be the opposite of the sign of the autocorrelation of the marginal rate of substitution. In the current model with $x_t$ i.i.d, Eq. (5a) implies that the sign of this autocorrelation is the opposite of the sign of $\theta$. Therefore, the term premium has the same sign as $\theta$. 


the price of this security in period \( t \). The gross rate of return on security \( L \) is \( RL_{t+1} = \xi_{t+1}/P_L \). In order to make security \( L \) correspond to (one-period) levered equity, specify the future payoff as \( \xi_{t+1} = \max[0, x_{t+1} - B]C_t \), where \( B \geq 0 \) is a constant. Observe that \( x_{t+1}C_t = C_{t+1} \) is per capita output in period \( t + 1 \), and \( B \) is interpreted as a fixed payment in period \( t + 1 \) promised to bondholders. Applying Eq. (9) along with the expression for the equilibrium marginal rate of substitution in Eq. (5a) yields

\[
E_t[R_{L_{t+1}}(B)] = \frac{E_t[\max[0, x_{t+1} - B]]}{\beta x_t^c E_t[\max[0, x_{t+1} - B]]}^{1/\gamma}.
\]

where the notation for the expected rate of return on the left-hand side emphasizes the dependence on \( B \).

5.2. An approximation to a one-period levered asset

Now I will use the canonical asset with \( n = 1 \) to approximate the expected return on security \( L \). Consider a one-period canonical asset that pays \( C_{t+1} = (x_{t+1}C_t) = (x_{t+1} - B)^{1/\gamma} \) in period \( t + 1 \). Under the assumed lognormality of \( x_{t+1} \), the coefficient of variation of \( (x_{t+1}C_t)^{1/\gamma} \) is \( \sqrt{\exp(\lambda^2 \sigma_x^2)} - 1 \). For security \( L \), the coefficient of variation of \( \xi_{t+1} \) is \( \sqrt{\text{Var}[\max[0, x - B]]/(E[\max[0, x - B]])} \). Therefore, the payoffs \( (x_{t+1}C_t)^{1/\gamma} \) and \( \xi_{t+1} \) will have equal coefficients of variation if \( \lambda = \lambda_1 \) where

\[
\lambda_1 = \frac{1}{\sigma_x} \ln \left( 1 + \frac{\text{Var}[\max[0, x - B]]}{(E[\max[0, x - B]])^2} \frac{1}{\lambda} \right).
\]

Even with the simplifying assumption of lognormality, the analytic expression for \( \lambda_1 \) in terms of the parameters of the distribution of \( x \) is cumbersome. This expression is relegated to Appendix A where it is also shown that for values of \( B \) that are small enough that the probability of default, \( \text{Pr}\{x < B\} \), is negligible, the coefficient of variation of \( \xi \) is approximately equal to the coefficient of variation of \( x^{1/\gamma} \) where

\[
\lambda_2 = \frac{1}{\lambda} \frac{B}{1 - E[x]}.
\]

For values of \( B \) small enough that the probability of default, \( \text{Pr}\{x < B\} \), is negligible, the introduction of \( B > 0 \) is essentially a leftward translation of the distribution of payoffs in which the mean payoff is reduced to a fraction \( 1 - B/E\{x\} \) of its original value and the standard deviation of the payoff is unchanged. This reduction in the mean payoff increases the coefficient of variation of the payoff to \( \lambda_2 \) times its value when \( B = 0 \).
Fig. 1. Values of $\lambda_1$ and $\lambda_2$.

Fig. 1 shows the values of $\lambda_1$ and $\lambda_2$ as functions of $B/E\{x\}$. The distributional assumption underlying Fig. 1 is that $\ln x$ is $N(0.018, (0.036)^2)$ where the mean and variance of the distribution were chosen to match the historical behavior of the annual growth rate of consumption per capita in the United States as reported by Mehra and Prescott (1985). For values of $B/E\{x\}$ small enough that $\Pr\{x < B\}$, the probability of default shown in Fig. 2, is close to zero, $\lambda_2$ is an excellent approximation for $\lambda_1$. As $B/E\{x\}$ increases toward 1.0, the probability of default increases noticeably and the values of $\lambda_1$ and $\lambda_2$ diverge somewhat.

To see how well the canonical asset approximates levered equity, I compute the expected rate of return on the one-period canonical asset, and compare this expected rate of return with that on one-period levered equity in Eq. (24). The rate of return on the one-period canonical asset is easily computed from Eq. (13) by setting $z \equiv x$ and using the lognormality of $x$ to obtain

$$E\{R_{t+1}(1, \lambda)\} = \frac{1}{\beta x_t} \exp \left[ A\mu + A\sigma^2 \left( \frac{1}{2} A \right) \right].$$

Fig. 2 shows the effect of leverage on the expected rate of return and assesses the quality of the approximation offered by the one-period canonical asset for
the case in which $A = 10$.\footnote{The calculation of the expected rate of return on security $L$ requires the calculation of moments of a truncated lognormal distribution. See Appendix A.} Because the purpose of the calculations reported here is to examine the role of leverage on expected rates of return, I have simply set $\beta x_i^\theta = 1$ in calculating the rates of return. Increasing $\beta x_i^\theta$ slightly would decrease the expected rates of return reported in Fig. 2, thereby making these rates look more realistic.\footnote{In fact, in Section 7, $\beta = 1.085$ and $\theta = 1.575$, so that evaluating $\beta x_i^\theta$ at the median value of $x_i = 1.018$ yields $\beta x_i^\theta = 1.12$, which would reduce the expected rates of return in Fig. 2 by about 12 percentage points.} Fig. 2 illustrates two important points. First, leverage can have a substantial effect on the expected rate of return even when the probability of default is essentially zero. In the absence of leverage ($B = 0$), the expected rate of return on security $L$ is 13.67\%, but when $B/E\{x\} = 0.6$, the expected rate of return on security $L$ increases to 15.94\%, even though the probability of default is smaller than $10^{-14}$. Second, Fig. 2 shows that the expected rates of return calculated using the canonical asset with $\lambda = \lambda_1$ or
\( \lambda = \lambda_2 \) are excellent approximations to the exact expected rate of return calculated from Eq. (24), provided that the probability of default is sufficiently small.

6. Moments of returns

In this Section I focus on the moments of returns on a one-period riskless asset and on levered equity. To simplify the resulting expressions, I assume that \( z_{t+1} \equiv x_{t+1} \) as is standard in applications of the Lucas fruit-tree model.\(^{12}\) This assumption implies that \( \mu_x = \mu_z = \mu \) and \( \sigma_x^2 = \sigma_{xz}^2 = \sigma_z^2 = \sigma^2 \). In addition I assume that \( a_j \equiv 1, j = 0, \ldots, n - 1 \), where \( n \) may be infinite.

The aim of this section is to understand how the unconditional means and variances of asset returns depend on the underlying preference parameters and payoff characteristics. I derive exact expressions for these moments of asset returns but some of the resulting expressions are too cumbersome to clearly reveal the effects of parameter values and payoffs on moments of returns. A common approach when facing such cumbersome analytic expressions is to calibrate the model and simulate it for various choices of parameter values. Rather than pursue a numerical strategy, I will pursue an analytical strategy by deriving first-order approximations that clearly illustrate the relationships between the underlying parameters and the moments of returns. It is worth emphasizing that I do not derive the solution to a problem in which the objective function and/or constraints have first been linearized.\(^{13}\) Instead, I first obtain exact solutions to the nonlinear problem, derive closed-form expressions for the moments of equilibrium returns, and then approximate the moments. The quality of these approximations is shown to be excellent in Table 1, which appears in Section 7.

6.1. Exact solutions

The unconditional riskless rate is calculated using Eq. (18) and the lognormality of \( x_{t+1} \) to obtain

\[
E[R_{t+1}(1, 0)] = \beta^{-1} \exp \left( (A - \theta) \mu - \frac{1}{2} (A^2 - \theta^2) \sigma^2 \right).
\] (28)

\(^{12}\) The Lucas fruit-tree model also restricts \( \lambda = 1 \). I do not impose this restriction here.

\(^{13}\) Campbell (1993) and Restoy and Weil (1995) (log)linearize the budget constraint of a representative consumer and use the approximate budget constraint to derive and analyze asset returns.
The unconditional expected rate of return on equity is calculated by setting \( n = \infty \) in Eqs. (12) and (14a), and then using Eqs. (21), (23) and (28) to obtain

\[
E[R_{t+1}(\infty, \lambda)] = \beta^{-1}[1 + \kappa(\lambda)(\exp(\theta A \sigma^2) - 1)][\exp(A \lambda \sigma^2)]
\]

\[
\times \exp\left(A - \theta\mu - \frac{1}{2}(A^2 - \theta^2)\sigma^2\right).
\]  

(29)

To get an explicit expression for the term premium, set \( a_i = 1, i = 0, 1, 2, \ldots \) in Eq. (16b) and use Eq. (21) to rewrite Eq. (16a) as

\[
TP_t(n, \lambda) = \frac{\kappa(\lambda) - [\kappa(\lambda)]^n}{1 - [\kappa(\lambda)]^n}\left[\exp(\theta A \sigma^2) - 1\right].
\]  

(30)

According to Eq. (30), the magnitude of the term premium increases monotonically in \( n \); the term premium is positive and monotonically increasing in \( n \) if \( \theta > 0 \) and is negative and monotonically decreasing in \( n \) if \( \theta < 0 \).

Exact expressions for the unconditional variances of rates of return are more complicated and are contained in Appendix B.

6.2. First-order approximations to moments of returns

In this section I present first-order approximations to moments of returns. Specifically, I treat the moments of returns as functions of \( \mu, \sigma^2, \) and \( \beta, \) and linearize the functions around \((\mu, \sigma^2, \beta) = (0, 0, 1)\).

The approximate unconditional mean of the riskless rate is calculated by linearizing Eq. (28) to obtain

\[
E'[R_{t+1}(1, 0)] \approx 2 - \beta + (A - \theta)\mu - \frac{1}{2}(A^2 - \theta^2)\sigma^2.
\]  

(31)

The approximate unconditional mean of the rate of return on equity is calculated by linearizing Eq. (29) to obtain\(^{14}\)

\[
E[R_{t+1}(\infty, \lambda)] \approx 2 - \beta + (A - \theta)\mu - \left[\frac{1}{2}A^2 - \frac{1}{2}\theta^2 - (\theta + \lambda)A\right]\sigma^2.
\]  

(32)

It follows immediately from Eqs. (31) and (32) that the first-order approximation to the unconditional equity premium \( E[R_{t+1}(\infty, \lambda) - R_{t+1}(1, 0)] \) is

\[
E[R_{t+1}(\infty, \lambda) - R_{t+1}(1, 0)] \approx (\theta + \lambda)A\sigma^2.
\]  

(33)

\(^{14}\) Note that \( \kappa(\lambda) = \beta = 1 \) at the point of linearization.
The approximate equity premium is the sum of the approximate risk premium \( \lambda A \sigma^2 \) (see Eq. (23)) and the approximate term premium\(^{15} \theta A \sigma^2 \).

The approximate unconditional variances of returns are calculated in Appendix B. The approximate unconditional variance of the riskless rate (see Eq. (B.20)) is

\[
\text{Var}\{R_{t+1}(1, 0)\} \approx \theta^2 \sigma^2. \tag{34}
\]

and the approximate unconditional variance of the rate of return on equity (see Eq. (B.22)) is

\[
\text{Var}\{R_{t+1}(\infty, \lambda)\} \approx [(\theta + \lambda)^2 + \theta^2] \sigma^2. \tag{35}
\]

### 7. Matching sample means and variance of rates of return

In this section I show how to choose values of the preference parameters \( A, \beta, \) and \( \theta, \) and the leverage parameter \( \lambda, \) so that the unconditional means and variances of the riskless rate and the rate of return on equity in the model match the sample values of these means and variances. The unconditional moments of returns depend on the four parameter values, \( A, \beta, \theta, \) and \( \lambda, \) and on the parameters of the lognormal distribution of the growth rate of consumption \( x_t. \) I will choose values of this lognormal distribution to match the sample moments of the growth rate of aggregate consumption per capita reported by Mehra and Prescott (1985) for annual data during the period 1889–1978. Specifically, I will set the mean growth rate, \( \mu, \) equal to 0.018 and the standard deviation of the growth rate, \( \sigma, \) equal to 0.036.

Let \( m_f \) and \( m_e \) represent the sample means of the riskless rate and the rate of return on levered equity, respectively. Similarly, let \( s_f \) and \( s_e \) represent the respective sample standard deviations of these rates of return. The values of these four sample moments, as reported by Mehra and Prescott (1985), are shown in the last column (labeled ‘Empirical values’) in Table 1. I will match the four sample moments with the corresponding approximate unconditional moments implied by the model using Eqs. (31), (33)–(35). Given the sample values of the four moments to be matched, this set of equations is a recursive system of four equations in four parameters to be determined. I will use a circumflex (\(^\wedge\)) to denote the values of the parameters that solve this system of equations.

First, choose \( \theta \) to match the variability of the riskless rate by setting the approximate unconditional variance of the riskless rate in Eq. (34) equal to \( s_f^2 \) and solving for \( \theta \) to obtain

\[
\bar{\theta} = \frac{s_f}{\sigma}. \tag{36}
\]

\(^{15}\)To calculate the term premium for levered equity, set \( n = \infty \) in Eq. (30) to obtain \( TP(\infty, \lambda) = \kappa(\lambda) \exp(\theta A \sigma^2) - 1. \) The approximate term premium is derived by linearizing this expression, noting that \( \kappa(\lambda) = 1 \) at the point of linearization, to obtain \( TP(\infty, \lambda) \approx \theta A \sigma^2. \)
Table 1
Matching the moments of returns ($\mu = 0.018$, $\sigma = 0.036$)

**Panel A: Parameters chosen to match sample moments**

<table>
<thead>
<tr>
<th>Parameter</th>
<th>Value</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\hat{\lambda}$</td>
<td>2.7411</td>
</tr>
<tr>
<td>$\hat{A}$</td>
<td>11.0483</td>
</tr>
<tr>
<td>$\hat{\theta}$</td>
<td>1.5750</td>
</tr>
<tr>
<td>$\hat{\beta}$</td>
<td>1.0850</td>
</tr>
</tbody>
</table>

**Panel B: Implied values of preference parameters (with 3 additional restrictions: $\gamma_0 = 0$, $\gamma_0 + \gamma_1 + \gamma_2 = 1$, $G = 1 + \mu$)**

<table>
<thead>
<tr>
<th>Parameter</th>
<th>Value</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\alpha$</td>
<td>11.0483</td>
</tr>
<tr>
<td>$\gamma_1$</td>
<td>0.1567</td>
</tr>
<tr>
<td>$\gamma_2$</td>
<td>0.8433</td>
</tr>
<tr>
<td>$\delta$</td>
<td>0.0720</td>
</tr>
<tr>
<td>$G$</td>
<td>1.018</td>
</tr>
</tbody>
</table>

**Panel C: Implied unconditional moments (percent per year)**

<table>
<thead>
<tr>
<th>Moment</th>
<th>Empirical values</th>
</tr>
</thead>
<tbody>
<tr>
<td>Riskless rate: mean</td>
<td>1.15 0.80</td>
</tr>
<tr>
<td>approx</td>
<td>0.80 0.80</td>
</tr>
<tr>
<td>Risky rate: mean</td>
<td>7.57 6.98</td>
</tr>
<tr>
<td>approx</td>
<td>6.98 6.98</td>
</tr>
<tr>
<td>Term premium</td>
<td>2.26 1.70</td>
</tr>
<tr>
<td>approx</td>
<td>2.26 1.70</td>
</tr>
<tr>
<td>Risk premium</td>
<td>4.00 (4.48)</td>
</tr>
<tr>
<td>approx</td>
<td>4.00 4.48</td>
</tr>
<tr>
<td>Riskless rate: standard deviation</td>
<td>5.74 5.67</td>
</tr>
<tr>
<td>approx</td>
<td>5.67 5.67</td>
</tr>
<tr>
<td>Risky rate: standard deviation</td>
<td>17.86 16.54</td>
</tr>
<tr>
<td>approx</td>
<td>16.54 16.54</td>
</tr>
</tbody>
</table>

**Note:** The period utility function is $u(c_t, v_t) = (c_t/v_t)^{1-\alpha}$, where $v_t \equiv C_t^{-\gamma_2}(G_t)^{\gamma_2}$, and the marginal rate of substitution is $M_{x,t+1}^{x} = \beta x_{t+1}^{x}/x_t^{x}$, where $\beta \equiv G^{\gamma_2(\gamma_2-1)/(1+\delta)} > 0$, $A \equiv x - \gamma_0(x - 1) > 0$, and $\theta \equiv \gamma_1(x - 1)$.

Second, choose the leverage parameter $\hat{\lambda}$ to match the approximate variability of the rate of return on levered equity by setting the approximate unconditional variance of this rate of return in Eq. (35) equal to $s_e^2$ and solving for $\hat{\lambda}$ to obtain:

$$\hat{\lambda} = \left[ \sqrt{\frac{s_e^2}{s_f^2}} - 1 \right] \hat{\theta}.$$  \hspace{1cm} (37)

\hspace{1cm} \hspace{1cm} 16 Note that $\hat{\lambda}$ will be real if and only if $s_e \geq s_f$, and $\hat{\lambda}$ will be positive if and only if $s_e \geq s_f \sqrt{2}$. Both of these conditions are satisfied in the sample used in this paper.
Third, choose the preference parameter $A$ to match the equity premium by setting the approximate unconditional mean of the equity premium in Eq. (33) equal to $m_e - m_f$ to obtain

$$
\hat{A} = \frac{m_e - m_f}{(\hat{\theta} + \hat{\lambda})\sigma^2}.
$$

(38)

Finally, choose the value of $\beta$ to match the average riskless rate by setting the approximate unconditional mean of the riskless rate in Eq. (31) equal to $m_f$ to obtain

$$
\hat{\beta} = 2 - m_f + (\hat{A} - \hat{\theta})\mu - \frac{1}{2}(\hat{A}^2 - \hat{\theta}^2)\sigma^2.
$$

(39)

Given the empirical values of $\mu, \sigma, m_e, m_f, s_e,$ and $s_f$, the values of $\hat{\theta}, \hat{\lambda}, \hat{A},$ and $\hat{\beta}$ implied by Eqs. (36)–(39) are shown in the top panel of Table 1. The leverage parameter that allows the model to match the sample moments of returns is $\hat{\lambda} = 2.74$. Interpreting $B/E_{t\{x_{t+1}'}$ as the leverage ratio, and setting $\lambda_2 = 2.74$ in Eq. (26) implies that the leverage ratio is 0.64, which is in the middle of the range reported by Masulis (1988) for leverage ratios based on book value, but is above the values of the leverage ratio based on market value.\textsuperscript{17} An alternative way to gauge the plausibility of $\hat{\lambda} = 2.74$ is to observe that dividends in the model are proportional to $y^{t}_j = c^{t}_j$, where $y_t$ is output per capita and $c_t$ is consumption per capita. Therefore, the (gross) growth rate of dividends is $x^{t}_j$ where $x_t$ is the growth rate of output and consumption, and the standard deviation of the growth rate of dividends is approximately $\lambda$ times the standard deviation of the growth rate of output and consumption.\textsuperscript{18} Using the standard deviations of growth rates reported in Cecchetti et al. (1990), (Table 1, p. 402), the standard deviation of the growth rate of dividends is 2.5 times the standard deviation of the growth rate of output, and 3.6 times the standard deviation of the growth rate of consumption. Thus the dividend variability implied by $\hat{\lambda} = 2.74$ seems consistent with the data.

The general formulation of preferences introduced in Section 2 has six parameters, but this set of parameters is under-identified. In addition to the moment conditions in Eqs. (36), (38) and (39), three restrictions are needed to obtain unique values of the six parameters. I will use the following three restrictions:

\textsuperscript{17} Masulis (1988), (Tables 1–3, pp. 8–9) reports leverage ratios for U.S. corporations. The leverage ratio based on market value ranged from 0.13 to 0.44 during the period 1929–1986, and the leverage ratio based on book value ranged from 0.53 to 0.75 during the period 1937–1984.

\textsuperscript{18} Since $\ln x_t$ is $N(\mu, \sigma^2)$, $\text{Var}(x^{t}_j) = \exp(2\lambda\mu + \lambda^2\sigma^2)[\exp(\lambda^2\sigma^2) - 1] \approx \lambda^2\sigma^2$. Therefore $\text{s.d.} \ (x^{t}_j) \approx \lambda \text{s.d.}(x_t)$.
\( \gamma_0 = 0, \gamma_0 + \gamma_1 + \gamma_2 = 1, \) and \( G = 1 + \mu. \) The values of the five nonzero preference parameters implied by these restrictions are reported in the second panel of Table 1. The coefficient of relative risk aversion \( \alpha \) is slightly above \( 11, \) which is a little higher than the upper bound of the range considered by Mehra and Prescott (1985), but is much smaller than values reported by Kandel and Stambaugh (1990) and Campbell and Cochrane (1994). The value of \( \gamma_1 \) is much smaller than the value of 1.0 used in Abel (1990) and allows the value of \( \theta \) to be relatively low even though \( \alpha \) exceeds 10. The relatively low value of \( \theta \) keeps the implied unconditional variance of the riskless rate from being too high, while the relatively high value of \( \alpha \) allows for a substantial equity premium. The rate of time preference \( \delta \) is about \( 7\% \) per year, even though \( \beta \) exceeds 1.08. Because \( G \) is greater than one and \( \gamma_2 \) is not negligible, this positive rate of time preference is consistent with the value of \( \beta \) greater than one.\(^\text{19}\) As noted by Benninga and Protopapadakis (1990) and Kocherlakota (1990), a value of \( \beta \) greater than one can improve the empirical fit of fruit-tree asset-pricing models, in particular by reducing the otherwise high implied value of the riskless rate. The introduction of \( \gamma_2 \) in this formulation offers a rationale for \( \beta > 1. \)

Panel C of Table 1 reports the unconditional moments of rates of return. Two values are presented for each moment implied by the model. The first value is computed from the exact solution, and the second value is computed from the first-order approximation. The first-order approximations are generally close to the exact values—in some cases, remarkably so. The approximate values of the means and standard deviations of the riskless rate and the rate of return on equity are equal to the corresponding sample moments by construction. The exact values of these moments are generally quite close to the sample values because the approximations are so close to the exact values.

The empirical column in Table 1 also contains an estimate of the average term premium. Over the period 1926–1993 the (arithmetic) average return on long-term government bonds was 5.4% per year compared to an average return on US Treasury bills of 3.7% per year,\(^\text{20}\) implying a term premium on fixed-income securities of 170 basis points per year. With the parameter values in Table 1, the model predicts a term premium on perpetual securities of 226 basis points per year which is 56 basis points higher than the sample average term premium on long-term government bonds. However, the implied term premium in Table 1 applies to an asset of \textit{infinite} maturity with \textit{risky} payoffs \( (\lambda = 2.74), \) whereas the empirical term premium is for long-term government bonds, which are assets of \textit{finite} maturity with \textit{known} payoffs \( (\lambda = 0). \) As it turns out, however,\(^\text{19}\) Although \( \beta > 1, \) the value of the utility function in Eq. (1) is finite. Recall from footnote 3 that the utility function will be finite if \( \beta E(x^{1-\delta + \theta}) < 1. \) For the parameters values in Table 1, \( \beta E(x^{1-\delta + \theta}) = 0.9759. \)

the quantitative impact of these differences in risk and maturity on the term premium is small. Reducing $\lambda$ from 2.74 to zero reduces the annual term premium on an infinite-maturity canonical asset by 5 basis points.\(^{21}\) The effect of maturity on the term premium depends on the ratio of the coupon to the face value. Reducing the maturity from infinite to 20 years has virtually no effect if the coupon-face ratio is 0.03, reduces the term premium by about 2 basis points if the coupon-face ratio is 0.06, and reduces the term premium by about 4 basis points if the coupon-face ratio is 0.09. Thus, a small part of the 56-basis point gap in Table 1 between the implied term premium on equity and the sample term premium on long-term government bonds is due to the finite maturity of bonds and to the variability of payoffs to equity. These effects imply that the gap between the term premium predicted by the model and the sample average term premium is smaller than 56 basis points per year.

Now consider the risk premium. The model predicts yields a risk premium of 4% per year. There is no direct empirical counterpart to the risk premium defined in Eq. (17). The value of 4.48% of the risk premium reported in parentheses in the final column of Table 1 is simply the empirical value of the equity premium minus the empirical value of the term premium.

The question that motivated the research in this paper was how much of the equity premium is a term premium and how much is a risk premium. According to the model, about one third of the equity premium is a term premium, and according to the empirical values in Table 1, the term premium accounts for about one fourth of the equity premium. Taken together, these results suggest that the risk premium is about 2 or 3 times as large as the term premium.

8. Concluding remarks

I have developed a simple model of asset pricing that can account for the unconditional means and variances of the rates of return on the riskless asset and on levered equity. In addition to providing closed-form expressions for the exact values of these moments, I have derived first-order approximations to these moments. These approximations permitted the development of an algorithm that allows the values of these unconditional moments in the model to match historical sample moments. In addition, these approximations provide a clear view of how the parameters of the model affect the moments of returns implied by the model. The approximate equity premium is the sum the approximate term premium $\theta A \sigma^2$ and the approximate risk premium $\lambda A \sigma^2$. The term premium and the risk premium are both proportional to $A \sigma^2$, where $A$ is

\(^{21}\)The first-order approximation for the term premium (see footnote 15) indicates that $\lambda$ does not have a first-order effect on the term premium.
The parameter $A$ is defined as $a!c_0$ ($a!1$). Under the assumption that $c_0'0$ as in Table 1, $A'a$ which is the coefficient of relative risk aversion.

A curvature parameter in the utility function and $\sigma$ is the standard deviation of consumption growth. However, there is an interesting difference in the factors that multiply $A\sigma^2$: in the case of the term premium, this factor is $\theta$ which depends only on preferences and is independent of the characteristics of the asset; in the case of the risk premium, this factor is $\lambda$ which is independent of preferences and is directly related to the volatility of the payoff on the asset.

The equity premium in this framework is $(\theta + \lambda)A\sigma^2$ compared with $A\sigma^2$ in a standard model without catching up with the Joneses preferences ($\theta = 0$) and without leverage ($\lambda = 1$). With $\theta = 1.58$ and $\lambda = 2.74$, the model used here yields an equity premium that is $\theta + \lambda = 4.32$ times as high as the equity premium in a standard model (where $\theta + \lambda = 1$) for given values of the coefficient of relative risk aversion $^22 A$ and consumption growth variance $\sigma^2$. Thus, the model can match the equity premium with a value of $A = 11.05$ rather $4.32 \times 11.05 = 47.74$. The increase in the equity premium relative to a standard model is due in about equal amounts to the change in preferences, which contributes $1.58 A\sigma^2$, and the increase in the variability of the payoff, which contributes $1.74 A\sigma^2$.

By using a high enough value of $\theta$, the term premium $\theta A\sigma^2$ could be made large enough to account for the empirical equity premium without appealing to leverage. However, such a high value of $\theta$ would make the implied riskless rate too volatile. The unconditional standard deviation of the riskless rate is $\theta \sigma$, so there is an intimate link between the term premium and the standard deviation of the short-term riskless rate: both are proportional to $\theta$. In economic terms, it is the volatility of the short-term riskless rate that gives rise to a term premium. If the riskless rate were constant over time, there would be no variability in the price of fixed-income securities, and there would be no term premium.

Because an increase in $\theta$ increases the volatility of the riskless rate, one cannot rely on catching up with the Joneses preferences alone to match both the expected equity premium and the volatility of the short-term riskless rate in an i.i.d. framework. Here is where leverage plays an important role. Increasing the value of the leverage parameter $\lambda$ directly increases the risk premium and hence the equity premium implied by the model, without increasing the implied volatility of the short-term riskless rate.

The formulation of leverage in this paper is perhaps the most novel modeling aspect, and further research along this line is needed. Although Section 5 demonstrates that $\lambda$ captures the effect of leverage on the expected return on a levered asset, this analysis is restricted to one-period assets and one-period senior claims against the payoffs to these assets. Additional research is needed to determine how best to model infinite-horizon assets that have issued multi-period fixed claims. This research would shed light on the role of leverage in

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$^22$ The parameter $A$ is defined as $a - \gamma_0(a - 1)$. Under the assumption that $\gamma_0 = 0$ as in Table 1, $A = a$ which is the coefficient of relative risk aversion.
asset pricing and also help to determine the appropriate empirical counterpart to be used to calibrate the degree of leverage in the model.

The analysis in this paper was confined to a fruit-tree framework in which equilibrium consumption per capita is an exogenous stochastic process. This framework is a useful initial testing ground for the asset pricing implications of leverage and catching up with the Joneses preferences. Since the model is capable of producing moments of returns that match the corresponding sample moments, the next step is to extend the model to allow the economy to transfer goods across time by capital investment. This richer framework will permit analysis of the implications for business cycles and growth as well as for asset returns.

Acknowledgements

I thank Allan Meltzer for correspondence that stimulated me to think about some of the issues addressed in this paper. I also thank John Cochrane, Urban Jermann, Alex Shapiro, Robert Stambaugh, Harald Uhlig, Philippe Weil, the participants in CEPR/Studienzentrum Gerzensee Conference on Macroeconomics and Finance, the Penn Macro Lunch Group, and seminars at Boston University, Duke University, Georgetown University, Michigan State University, the University of Chicago, the University of Michigan, and the Asset Pricing Program of the National Bureau of Economic Research for helpful comments. Research support from the National Science Foundation is gratefully acknowledged.

Appendix A. Moments of truncated lognormal distributions

Proposition 1. If \( x = \ln y \) is \( N(\mu, \sigma^2) \) and \( B > 0 \), then \( E\{y^a|y > B\} = \frac{1}{\rho} \left[ 1 - \Phi\left( \ln B - (\mu + a\sigma^2) \right) \right] E\{y^a\} \) where \( \rho \equiv 1 - \Phi\left( \frac{\ln B - \mu}{\sigma} \right) \) and \( \Phi() \) is the standard normal c.d.f.

Proof. Note that \( e^x = y \Rightarrow e^{ax} = y^a \) and \( y \geq B \Rightarrow \ln y \geq \ln B \Rightarrow x \geq \ln B \). Therefore,

\[
E\{y^a|y > B\} = \frac{1}{\rho} \int_{\ln B}^{\infty} \frac{1}{\sigma \sqrt{2\pi}} \exp\left( -\frac{1}{2\sigma^2}(x - \mu)^2 \right) \exp(ax) \, dx. \tag{A.1}
\]
Combining the terms to be exponentiated, completing the square, and rearranging yields

\[ E\{y^a | y > B\} \]

\[ = \frac{1}{\rho} \exp \left( \mu a + \frac{1}{2} a^2 \sigma^2 \right) \frac{1}{\sigma \sqrt{2\pi}} \int_{\ln B}^{\infty} \exp \left( -\frac{1}{2\sigma^2} (x - (\mu + a\sigma^2))^2 \right) dx. \]

(A.2)

Recognizing that \( E\{y^a\} = \exp(\mu a + (1/2)a^2\sigma^2) \) yields

\[ E\{y^a | y > B\} = \frac{1}{\rho} \left[ 1 - \Phi \left( \frac{\ln B - (\mu + a\sigma^2)}{\sigma} \right) \right] E\{y^a\}. \]

(A.3)

**Corollary.**

\[ E\{y^{-A}\max[0, y - B]\} \]

\[ = \left[ 1 - \Phi \left( \frac{\ln B - (\mu + (1 - A)\sigma^2)}{\sigma} \right) \right] E\{y^{1-A}\} - \]

\[ B \left[ 1 - \Phi \left( \frac{\ln B - (\mu - A\sigma^2)}{\sigma} \right) \right] E\{y^{-A}\}. \]

**Proposition 2:** Let \( \xi \equiv \max[y - B, 0] \) where \( x \equiv \ln y \) is \( N(\mu, \sigma^2) \). Then

\[ \frac{\text{Var}\{\xi\}}{(E\{\xi\})^2} = \frac{E\{y^2 | y > B\} - (E\{y | y > B\})^2}{\rho(E\{y | y > B\} - B)^2} + \frac{1}{\rho} - 1, \]

where \( \rho \equiv 1 - \Phi \left( \frac{\ln B - \mu}{\sigma} \right) \).

**Proof.**

\[ \frac{\text{Var}\{\xi\}}{(E\{\xi\})^2} = \frac{E\{\xi^2\} - (E\{\xi\})^2}{(E\{\xi\})^2} = \frac{E\{\xi^2\}}{(E\{\xi\})^2} - 1, \]

(A.4)

\[ E\{\xi\} = E\{\max[0, y - B]\} = \rho(E\{y | y > B\} - B), \]

(A.5)

\[ (E\{\xi\})^2 = \rho^2((E\{y | y > B\})^2 - 2BE\{y | y > B\} + B^2), \]

(A.6)

\[ E\{\xi^2\} = E\{\{\max[0, y - B]\}^2\} = \rho E\{[y - B]^2 | y > B\} \]

\[ = \rho(E\{y^2 | y > B\} - 2BE\{y | y > B\} + B^2), \]

(A.7)

\[ E\{\xi^2\} = \rho(E\{y^2 | y > B\} - (E\{y | y > B\})^2) + \frac{1}{\rho} (E\{\xi\})^2. \]

(A.8)
Substituting Eqs. (A.6) and (A.8) into Eq. (A.4) and using Eq. (A.5) yields

$$\frac{\text{Var}\{\xi\}}{(E\{\xi\})^2} = \frac{\{y^2|y > B\} - (E\{y|y > B\})^2}{\rho(E\{y|y > B\} - B)^2} + \frac{1}{\rho} - 1. \quad (A.9)$$

**Corollary.** If $\rho$ is approximately equal to 1, then

$$\frac{\text{Var}\{\xi\}}{(E\{\xi\})^2} \approx \frac{1}{(1 - \frac{B}{E\{y\}})^2} \frac{\text{Var}\{y\}}{(E\{y\})^2}.$$  

**Proof.** For $\rho$ sufficiently close to one, realizations of $y < B$ are of negligible importance. Therefore, Proposition 2 implies

$$\frac{\text{Var}\{\xi\}}{(E\{\xi\})^2} \approx \frac{E\{y^2\} - (E\{y\})^2}{(E\{y\} - B)^2} = \frac{E\{y^2\}}{(E\{y\})^2} - \frac{1}{(1 - \frac{B}{E\{y\}})^2} \frac{E\{y^2\} - (E\{y\})^2}{(E\{y\})^2} = \frac{1}{(1 - \frac{B}{E\{y\}})^2} \frac{\text{Var}\{y\}}{(E\{y\})^2}.$$  

**Appendix B. Moments of rates of return**

This appendix calculates the first and second moments of rates of return on riskless bills and levered equity under the assumption that $\ln z_{t+1} \equiv \ln x_{t+1}$ is $N(\mu, \sigma^2)$. For equity, $a_j = 1$ for $j = 0, \ldots, n - 1$, so $\omega(n, \lambda) = \omega(1, \lambda)(1 - \kappa^n)/(1 - \kappa)$. Using this expression for $\omega(n, \lambda)$ in Eq. (8), the realized rate of return on the canonical asset can be written as

$$R_{t+1}(n, \lambda) = x_t^{-\theta}H(x_{t+1}; n, \lambda), \quad (B.1)$$

where

$$H(x_{t+1}; n, \lambda) \equiv \left( \frac{1}{1 - \kappa^n} \right) \left[ (1 - \kappa^{n-1})x_{t+1}^{\theta+\lambda} + \frac{1 - \kappa}{\omega(1, \lambda)} x_{t+1}^\lambda \right] \quad (B.2)$$

and $\kappa(\lambda)$ is written simply as $\kappa$.  

Using the fact that $\ln x_{t+1}$ is $N(\mu, \sigma^2)$ it can be shown that

$$
E\{H(x_{t+1}; n, \lambda)\} \equiv \left(\frac{1}{1 - \kappa}\right) \left((1 - \kappa^{n-1}) \exp\left((\theta + \lambda)\mu + \frac{1}{2} (\theta + \lambda)^2 \sigma^2\right) + \frac{1 - \kappa}{\omega(1, \lambda)} \exp\left(\lambda \mu + \frac{1}{2} \lambda^2 \sigma^2\right)\right]
$$

(B.3)

and

$$
\text{Var}\{H(x_{t+1}; n, \lambda)\}
= \left(\frac{1}{1 - \kappa}\right)^2 \left((1 - \kappa^{n-1})^2 \exp(2(\theta + \lambda)\mu)
+ (\theta + \lambda)^2 \sigma^2) \exp((\theta + \lambda)^2 \sigma^2) - 1\right)
+ \left(\frac{1 - \kappa}{\omega(1, \lambda)}\right)^2
\times \exp(2 \lambda \mu + \lambda^2 \sigma^2) \exp((\lambda^2 \sigma^2) - 1)
+ 2 \frac{(1 - \kappa^{n-1})(1 - \kappa)}{\omega(1, \lambda)}
\times \exp\left((\theta + 2 \lambda)\mu + \frac{1}{2} ((\theta + \lambda)^2 + \lambda^2) \sigma^2\right) \exp((\theta + \lambda) \lambda \sigma^2) - 1\right]
$$

(B.4)

For riskless one-period bills, $n = 1$ and $\lambda = 0$. The first and second moments of $H(x_{t+1}; n, \lambda)$ are

$$
E\{H(x_{t+1}; 1, 0)\} = \frac{1}{\omega(1, 0)}
$$

(B.5)

and

$$
\text{Var}\{H(x_{t+1}; 1, 0)\} = 0.
$$

(B.6)

For levered equity, $n = \infty$ and the first and second moments of $H(x_{t+1}; n, \lambda)$ are

$$
E\{H(x_{t+1}; \infty, \lambda)\} \equiv \exp\left((\theta + \lambda)\mu + \frac{1}{2} (\theta + \lambda)^2 \sigma^2\right)
+ \frac{1 - \kappa}{\omega(1, \lambda)} \exp\left(\lambda \mu + \frac{1}{2} \lambda^2 \sigma^2\right)
$$

(B.7)
and

$$\text{Var}\{H(x_{t+1}; \infty, \lambda)\}$$

$$= \exp(2(\theta + \lambda)\mu + (\theta + \lambda)^2\sigma^2)(\exp((\theta + \lambda)^2\sigma^2) - 1)$$

$$+ \left(\frac{1 - \kappa}{\omega(1, \lambda)}\right)^2 \exp(2\lambda\mu + \lambda^2\sigma^2)(\exp(\lambda^2\sigma^2) - 1)$$

$$+ 2\left(\frac{1 - \kappa}{\omega(1, \lambda)}\right) \exp((\theta + 2\lambda)\mu + \frac{1}{2}((\theta + \lambda)^2 + \lambda^2)\sigma^2)$$

$$\times (\exp((\theta + \lambda)\lambda\sigma^2) - 1).$$

(B.8)

**Unconditional moments.** The assumption of i.i.d. growth rates implies that $x_{t}^{-\theta}$ and $H(x_{t+1}; n, \lambda)$ are independent. I will use the following identities that hold for independent random variables $u$ and $v$:

$$E\{uv\} = E\{u\}E\{v\},$$

(B.9)

$$\text{Var}\{uv\} = \text{Var}\{u\}\text{Var}\{v\} + [E\{u\}]^2\text{Var}\{v\} + [E\{v\}]^2\text{Var}\{u\}.$$  

(B.10)

Therefore

$$E\{R_{t+1}(n, \lambda)\} = E\{x_{t}^{-\theta}\}E\{H(x_{t+1}; n, \lambda)\}$$

(B.11)

and

$$\text{Var}\{R_{t+1}(n, \lambda)\} = \text{Var}\{x_{t}^{-\theta}\}\text{Var}\{H(x_{t+1}; n, \lambda)\}$$

$$+ [E\{x_{t}^{-\theta}\}]^2\text{Var}\{H(x_{t+1}; n, \lambda)\}$$

$$+ [E\{H(x_{t+1}; n, \lambda)\}]^2\text{Var}\{x_{t}^{-\theta}\}. ($$)

(B.12)

The mean and variance of $H(x_{t+1}; n, \lambda)$ have already been calculated in Eqs. (B.2) and (B.4). The mean and variance of $x_{t}^{-\theta}$ are

$$E\{x_{t}^{-\theta}\} = \exp\left(-\theta\mu + \frac{1}{2}\theta^2\sigma^2\right) \approx 1 - \theta\mu + \frac{1}{2}\theta^2\sigma^2$$

(B.13)

and

$$\text{Var}\{x_{t}^{-\theta}\} = \exp( - 2\theta\mu + \theta^2\sigma^2)[\exp(\theta^2\sigma^2) - 1] \approx \theta^2\sigma^2.$$  

(B.14)
**Linearization.** Recall that \( a_0 = 1 \) so that under conditional lognormality Eq. (11) implies

\[
1/\omega(1, \lambda) = \beta^{-1} \exp \left( - (\lambda - A) \mu - \frac{1}{2} (\lambda - A)^2 \sigma^2 \right). \tag{B.15}
\]

Linearizing Eq. (B.15) around the point \((\mu, \sigma^2, \beta) = (0, 0, 1)\) yields

\[
\frac{1}{\omega(1, \lambda)} \approx 2 - \beta - (\lambda - A) \mu - \frac{1}{2} (\lambda - A)^2 \sigma^2. \tag{B.16}
\]

Under normality with \( x_{t+1} \equiv z_{t+1} \), Eq. (10b) implies

\[
\kappa = \beta \exp \left( (\theta + \lambda - A) \mu + \frac{1}{2} (\theta + \lambda - A)^2 \sigma^2 \right). \tag{B.17}
\]

Linearizing Eq. (B.17) around \((\mu, \sigma^2, \beta) = (0, 0, 1)\) yields

\[
\kappa \approx \beta + (\theta + \lambda - A) \mu + \frac{1}{2} \sigma^2 (\theta + \lambda - A)^2. \tag{B.18}
\]

The approximate unconditional mean of the riskless rate is calculated using the identity in Eq. (B.11) along with Eqs. (B.5) and (B.16) with \( \lambda = 0 \), and Eq. (B.13) to obtain

\[
E\{ R_{t+1}(1, 0) \} \approx 2 - \beta + (A - \theta) \mu + \frac{1}{2} (\theta^2 - A^2) \sigma^2. \tag{B.19}
\]

The approximate unconditional variance of the riskless rate is calculated using the identity in Eq. (B.12) along with Eqs. (B.6) and (B.14) to obtain

\[
\text{Var}\{ R_{t+1}(1, 0) \} \approx \theta^2 \sigma^2. \tag{B.20}
\]

The approximate unconditional mean of the rate of return on equity is calculated by substituting Eqs. (B.13) and (B.7) into Eq. (B.11), then using Eqs. (B.15) and (B.17) and linearizing to obtain

\[
E\{ R_{t+1}(\infty, \lambda) \} = 2 - \beta + (A - \theta) \mu + \frac{1}{2} (\theta^2 - A^2) \sigma^2 + (\theta + \lambda) A \sigma^2. \tag{B.21}
\]

The approximate unconditional variance of the rate of return on equity is calculated by substituting Eqs. (B.7), (B.8), (B.13) and (B.14), into Eq. (B.12), and linearizing to obtain

\[
\text{Var}\{ R_{t+1}(\infty, \lambda) \} \approx [(\theta + \lambda)^2 + \theta^2] \sigma^2. \tag{B.22}
\]
References