Journal of Banking & Finance 35 (2011) 67–81

Contents lists available at ScienceDirect

Journal of Banking & Finance

journal homepage: www.elsevier.com/locate/jbf

Future labor income growth and the cross-section of equity returns

Dongcheol Kim [a], Tong Suk Kim [b], Byoung-Kyu Min [c,*]

[a] Business School, Korea University, Seoul, Republic of Korea
[b] Graduate School of Finance, Korea Advanced Institute of Science and Technology (KAIST), Seoul, Republic of Korea
[c] Fisher College of Business, Ohio State University, OH 43204, USA

ARTICLE INFO

Article history:
Received 6 February 2010
Accepted 12 July 2010
Available online 15 July 2010

JEL classification:
G12
G14

Keywords:
Future labor income growth
Fama–French factors
Economic tracking portfolio
Intertemporal CAPM

ABSTRACT

This paper examines the equilibrium relation between future labor income growth and expected asset returns; it proposes revisions in the expectation of future labor income growth as a macroeconomic state variable and suggests a three-factor model, including a factor related to this variable, along with the consumption growth factor and the market factor. The proposed future labor income growth factor is positively associated with the Fama–French factors and subsumes their explanatory power in explaining the cross-section of stock returns. These results provide a possible economic explanation for the roles of the Fama–French factors: they are compensation for higher exposure to the risk related to changes in the value of human capital. This paper also compares the performance of the proposed three-factor model with other competing models and finds that the proposed model specification better captures cross-sectional variation in average returns than any of the competing asset pricing models considered.

1. Introduction

It is a stylized empirical fact in the literature that small stocks and value stocks have higher average returns than big stocks and growth stocks, respectively. The Capital Asset Pricing Model (CAPM) of Sharpe (1964) and Linter (1965) encounters difficulty in accounting for these well-established empirical regularities (Fama and French, 1992). In response to this difficulty, Fama and French (1993, 1995, 1996) propose a three-factor model that includes a factor related to size (SMB) and a factor related to book-to-market ratio (HML), together with the market factor. These authors empirically demonstrate that their model largely explains a cross-sectional pattern of average stock returns of portfolios sorted by size and book-to-market ratio.

Nonetheless the Fama–French model is often criticized, because its factors lack theoretical justification. Furthermore, its factors and test portfolios share the same characteristics. These considerations result in frequent debate over the interpretation of the success of the Fama–French factors. Fama and French (1993, 1995, 1996) argue that SMB and HML might mimic state variables of special hedging concern to investors; however, they have not identified which state variables SMB and HML proxy for. Specifically, Fama and French (1992, p. 450) suggest a path toward the economic meaning of their factors by stating that "examining relations between the returns on these portfolios and economic variables that measure variations in business conditions might help expose the nature of the economic risks captured by size and book-to-market equity."

Several studies have examined a link between the Fama–French factors and macroeconomic variables related to business cycle fluctuations. Liew and Vassalou (2000) show that SMB and HML have the ability to predict future economic growth. Vassalou (2003) argues that changes in the investment opportunity set are summarized by changes in future GDP growth, and SMB and HML appear to contain news mainly related to future GDP growth. Petkova (2006) and Hahn and Lee (2006) show that shocks to conditioning variables such as dividend yield, term spread, default spread, and one-month Treasury bill yield, which forecast future investment opportunity sets, fully replace the explanatory power of HML and SMB in the cross-section of average returns.

In this paper, we propose revisions in the expectation of future labor income growth as just such a macroeconomic state variable that is closely related to macroeconomic conditions and business cycle fluctuations and that may reveal the nature of the economic risk captured by size and book-to-market equity. We then suggest a three-factor model that includes a factor related to this variable, along with the consumption growth factor and the market factor. We examine whether revisions in the expectation of future labor income growth capture the pricing abilities of the Fama–French factors in explaining the size and book-to-market effects. In order

* Corresponding author. Tel.: +1 614 556 6694; fax: +1 614 292 2418.
E-mail addresses: kimdc@korea.ac.kr (D. Kim), tskim@business.kaist.ac.kr (T.S. Kim), min_104@fisher.osu.edu (B.-K. Min).

0378-4266/$ - see front matter © 2010 Elsevier B.V. All rights reserved.
doi:10.1016/j.jbankfin.2010.07.014

to obtain the risk factor that captures revisions in future labor income growth, which is unobservable, we adopt the economic tracking portfolio approach introduced by Lamont (2001). Economic tracking portfolios are designed to capture unexpected returns that are maximally correlated with unexpected components (or news) of a target macroeconomic variable (in this study, the discounted sum of future labor income growth).

We choose revisions in the expectation of future labor income growth as both a source of risk and a state variable of investors' hedging concerns for the following reasons: First, since investors fear to have low stock returns in bad times, when expectation of future labor income (or return on human capital) is changed to be low, stocks having positive correlation with news about future labor income would demand a high risk premium. Second, shocks to human capital are aggregate risks that affect the total wealth of a representative agent (Campbell, 1996). Rather than focusing on changes in current labor income growth as Jagannathan and Wang (1996) did, we argue that changes in labor income growth expectation are more important determination of the return on human capital.

We find that our proposed three-factor specification explains relatively well the cross-section of average returns for size and book-to-market sorted portfolios, and performs better than the competing asset pricing models considered: the Fama and French (1993) model, the CAPM, the Jagannathan and Wang (1996) human capital CAPM, the consumption CAPM of Breeden (1979), the Epstein–Zin (1991) model, the Lettau–Ludvigson (2001) model, and the Vassalou (2003) model. More importantly, the risk factor related with revisions in the expectation of future labor income growth is positively associated with the Fama–French factors SMB and HML, and subsumes the explanatory power of these Fama–French factors in explaining the cross-section of stock returns. Since the results could be sensitive to the specification used for constructing the tracking portfolio, we perform robustness tests using various alternative specifications for constructing the tracking portfolio. Nevertheless, the overall results are qualitatively the same.

Given the outcome in Vassalou (2003) showing a strong association of the future GDP growth factor with the Fama and French factors, what remains to be investigated is how our future labor income growth factor fares with the Vassalou's (2003) findings. Even more so is the fact that two macroeconomic variables are related with each other and the procedure of constructing both factors is similar. We find that the significance of the future GDP growth factor disappears in the presence of the future labor growth factor, while the significance of the future labor growth factor is still maintained. We also find that the remaining GDP growth component after excluding the labor income component loses its statistical significance and the association with the Fama and French factors. Our results imply that the Vassalou's (2003) finding might be driven by the relation between the future labor income growth and the Fama and French factors.

The positive association between the Fama–French factors and the labor income risk could be attributable to the asymmetry of employment across firms. In recession, employment in small and value firms, in which cash flows are uncertain and earnings are persistently low (Chan and Chen, 1991; Fama and French, 1995), is more vulnerable than in big and growth firms. Since small and value firms have high risk exposure to SMB and HML, a negative shock to SMB and HML may imply a negative shock to the value of human capital. Rational investors, who have hedging concern with respect to the state variable associated with human capital, have an incentive to avoid stocks of small and value firms. As a result, small and value firms are riskier than big and growth firms in recession, when the price of risk associated with labor income is high.

The rest of the paper is organized as follows: Section 2 explains the theoretical background of our three-factor model; Section 3 explains empirical methodology and data; Section 4 presents the empirical results; Section 5 reports various robustness tests; and Section 6 concludes.

2. Theoretical background

Since time-varying expectation of future labor income in the economy should capture movements in a relevant state variable, such as the level of human capital, it is likely to have an influence on equilibrium asset returns.[1] To see this formally, consider a representative agent whose utility is assumed to take the recursive form of Epstein and Zin (1989, 1991)

$$U_t = \left\{(1-\delta)C_t^{\frac{\psi-1}{\psi}} + \delta\left[E_t\left(U_{t+1}^{1-\gamma}\right)\right]^{\frac{1}{\theta}}\right\}^{\frac{\psi}{\psi-1}}, \tag{1}$$

where C_t is the consumption level at time t, $\theta \equiv (1-\gamma)/(1-\frac{1}{\psi})$, $\gamma > 0$ is the relative risk aversion coefficient, $\psi > 0$ denotes the elasticity of intertemporal substitution (EIS), and $0 < \delta < 1$ is the time discount factor. When $\theta = 1$, this reduces to the standard model of the time-separable power utility model.

The intertemporal budget constraint for a representative agent can be written as

$$W_{t+1} = R_{w,t+1}(W_t - C_t), \tag{2}$$

where W_{t+1} is the representative agent's total wealth and $R_{w,t+1}$ is the return on W_{t+1}. The representative agent's total wealth includes human capital as well as financial assets. From Eqs. (1) and (2), a Euler equation for asset i is obtained:

$$E_t\left\{\delta^\theta\left(\frac{C_{t+1}}{C_t}\right)^{-\frac{\theta}{\psi}}\left(\frac{1}{R_{w,t+1}}\right)^{1-\theta}R_{i,t+1}\right\} = 1. \tag{3}$$

Thus, the log stochastic discount factor or pricing kernel is equal to

$$m_{t+1} = \theta\log\delta - \frac{\theta}{\psi}\Delta c_{t+1} - (1-\theta)r_{w,t+1}, \tag{4}$$

where $\Delta c_{t+1} \equiv \log\left(\frac{C_{t+1}}{C_t}\right)$ and $r_{w,t+1} \equiv \log(R_{w,t+1})$ denote the log consumption growth and log return on total wealth, respectively.

When investors' total wealth consists of financial wealth and human capital, the aggregate return on total wealth can be expressed as

$$R_{w,t+1} = (1-\upsilon)R_{a,t+1} + \upsilon R_{h,t+1}, \tag{5}$$

where υ is the ratio of human capital wealth to total wealth, $R_{a,t+1}$ is the return on financial wealth, and $R_{h,t+1}$ is the return on human capital. Campbell (1996) shows that Eq. (5) can be approximated as the log or continuously compounded return:

$$r_{w,t+1} \approx (1-\upsilon)r_{a,t+1} + \upsilon r_{h,t+1}, \tag{6}$$

where $r_{t+1} = \log(R_{t+1})$.

In fact, labor income (Y_{t+1}) can be thought of as the dividend on human capital (H_{t+1}) (Campbell, 1996; Jagannathan and Wang, 1996). Thus, return on human capital ($R_{h,t+1}$) can be defined as

$$R_{h,t+1} = \frac{H_{t+1} + Y_{t+1}}{H_t}. \tag{7}$$

If we follow the log-linear approximation of Campbell and Shiller (1988) under the assumption of a constant discount rate on human capital following Shiller (1993), log human capital (h_t) can be ex-

[1] Pantzalis and Park (2009) document that their stock market valuation measure of human capital can predict future return performance.

pressed as a function of the discounted sum of future labor income growth (y_t)

$$h_t = h + y_t + E_t \sum_{j=0}^{\infty} \rho^j \Delta y_{t+1+j}, \tag{8}$$

where h is a constant of no interest.[2] This shows that the expected discounted value of labor income is an important determination of human capital wealth. Thus, the expectation of future labor income growth should contain important information about any state variable associated with human capital.

The log return on human capital ($r_{h,t+1}$) can be written by a linear combination of future log labor income growth:

$$r_{h,t+1} = r + (E_{t+1} - E_t) \sum_{j=0}^{\infty} \rho^j \Delta y_{t+1+j}, \tag{9}$$

where r is a constant of no interest.[3] That is, the return on human capital is determined by revision to expectations of future labor income growth. The last term on the right-hand side of Eq. (9) measures the contribution of news about future labor income growth to state variable h_t, and therefore captures the expected long-run wealth effect of labor income shocks.

Substituting Eq. (9) into Eq. (6) yields

$$r_{w,t+1} = r\upsilon + (1-\upsilon) r_{a,t+1} + \upsilon (E_{t+1} - E_t) \sum_{j=0}^{\infty} \rho^j \Delta y_{t+1+j}. \tag{10}$$

Then, Eq. (10) is substituted into Eq. (4) to obtain

$$m_{t+1} = k - \frac{\theta}{\psi} \Delta c_{t+1} - (1-\theta)(1-\upsilon) r_{a,t+1} - (1-\theta)\upsilon (E_{t+1} - E_t) \sum_{j=0}^{\infty} \rho^j \Delta y_{t+1+j}. \tag{11}$$

Eq. (11) indicates that the (log) stochastic discount factor depends on news about future long-horizon labor income growth, which is the last term on the right-hand side. Therefore, revisions in the expectation of future labor income growth should appear as an additional risk factor along with current consumption growth and current return on financial wealth.

For the cross-section of asset returns, the following expected return-covariance representation must hold in equilibrium:

$$E(r_{i,t+1} - r_{f,t+1}) + \frac{\sigma_i^2}{2} = -\mathrm{Cov}(m_{t+1}, r_{i,t+1}), \tag{12}$$

where $\sigma_i^2/2$ is a Jensen Inequality adjustment arising from the lognormal model, and the left-hand side of Eq. (12) is the relevant measure for risk premium for asset i. Substituting the pricing kernel equation of (11) into Eq. (12), we see that the expected risk premium for any asset i is determined by three covariances. That is,

$$E(r_{i,t+1} - r_{f,t+1}) + \frac{\sigma_i^2}{2} = \frac{\theta}{\psi}\sigma_{ic} + (1-\theta)(1-\upsilon)\sigma_{ia} + (1-\theta)\upsilon\sigma_{ih}, \tag{13}$$

where $\sigma_{ic} \equiv \mathrm{Cov}(r_{i,t+1}, \Delta c_{t+1})$, $\sigma_{ia} \equiv \mathrm{Cov}(r_{i,t+1}, r_{a,t+1})$, and $\sigma_{ih} \equiv \mathrm{Cov}\left[r_{i,t+1}, (E_{t+1} - E_t)\sum_{j=0}^{\infty}\rho^j \Delta y_{t+1+j}\right]$ represents the covariances of stock i's return with the current consumption growth (Δc_{t+1}), current return on financial wealth ($r_{a,t+1}$), and revision in the expectation of future labor income growth ($(E_{t+1} - E_t)\sum_{j=0}^{\infty}\rho^j \Delta y_{t+1+j}$), respectively.[4]

Given that most asset pricing models are estimated and evaluated in the form of an expected return-beta representation, we can restate Eq. (13) in terms of betas as

$$E(r_{i,t+1} - r_{f,t+1}) + \frac{\sigma_i^2}{2} = \lambda_c \beta_{ic} + \lambda_a \beta_{ia} + \lambda_h \beta_{ih}, \tag{14}$$

where $\lambda_c \equiv \frac{\theta}{\psi}\sigma_c^2$, $\lambda_a \equiv (1-\theta)(1-\upsilon)\sigma_a^2$, and $\lambda_h \equiv (1-\theta)\upsilon\sigma_h^2$ are the prices of risk for the three risk factors. Following Campbell and Vuolteenaho (2004), if we use simple expected returns, $E(R_{i,t+1} - R_{f,t+1})$, instead of log returns, $E(r_{i,t+1} - r_{f,t+1})$, Eq. (14) becomes

$$E(R_{i,t+1} - R_{f,t+1}) \approx \lambda_c \beta_{ic} + \lambda_a \beta_{ia} + \lambda_h \beta_{ih}, \tag{15}$$

where R indicates holding period return. Eq. (15) is the three-factor model that we put forth in this paper. The third component in Eq. (15), β_{ih}, indicates the risk exposure on the factor that reflects a revision in expectation of future labor income growth. It is notable that since the covariance terms in the above equations are figured separately, it is consistent with the theoretical derivation to estimate betas in Eq. (15) in separate univariate regression models.[5]

3. Empirical methodology and data

3.1. Construction of the economic tracking portfolio

In the three-factor model specification of Eq. (15), the first two risk factors (current consumption growth and current return on financial assets) are empirically well specified in the literature and so easily obtained. However, the third risk factor (revisions in the expectation of future labor income growth) is not. In order to obtain the risk factor that captures revisions in the expectation of future labor income growth, we adopt the economic tracking portfolio approach, which was introduced by Lamont (2001). Economic tracking portfolios are designed to capture unexpected returns that are maximally correlated with unexpected components (or news) of a target macroeconomic variable (in this study, the discounted sum of future labor income growth). The first assumption in this approach is that one can always write a projection equation of news on unexpected returns. That is,

$$(E_{t+1} - E_t) \sum_{j=0}^{\infty} \rho^j \Delta y_{t+1+j} = a\tilde{R}_{t,t+1} + \eta_{t+1}, \tag{16}$$

where $\tilde{R}_{t,t+1}$ is a vector of unexpected returns on the base assets, which are actual return minus expected return [$=R_{t,t+1} - E_t(R_{t,t+1})$], and η_{t+1} is the component of revisions or news that is orthogonal to unexpected returns.

The realization of labor income growth in all future periods, $\sum_{j=0}^{\infty}\rho^j \Delta y_{t+1+j}$, can be rewritten as

---

[2] Specifically, $h = \frac{k_h}{1-\rho_h}$, where $\rho_h \equiv \frac{1}{1+e^{\overline{y-h}}}$, and $k_h \equiv -\log\rho_h - (1-\rho_h)\log\left(\frac{1}{\rho_h} - 1\right)$.

[3] Specifically, $r = E_t r_{h,t+1}$.

[4] The last term on the right-hand side in Eq. (13) indicates that our model reflects changes in expectations of discounted labor income growth.

[5] This paper differs from Campbell (1996) in that consumption is not substituted out using the intertemporal budget constraint combined with the assumption of homoskedasticity for both asset returns and consumption growth. Since the model-implied consumption innovations, which are determined by news about current returns and by news about future expected returns on the market portfolio, differ entirely from those in the data (Lustig and Van Nieuwerburgh, 2008), we avoid this possible model misspecification error. More importantly, the main purpose of this paper is to investigate the relation between Fama–French factors and the state variable associated with human capital. Our model also differs from Jagannathan and Wang (1996). These authors assume that labor income growth is unpredictable. As a result, labor income growth over the following quarter becomes a risk factor. As documented in Campbell (1996) and our results, however, labor income growth is predictable in the data.

$$\sum_{j=0}^{\infty}\rho^j\Delta y_{t+1+j} = E_{t+1}\sum_{j=0}^{\infty}\rho^j\Delta y_{t+1+j} + e_{t+1}$$
$$= E_t\sum_{j=0}^{\infty}\rho^j\Delta y_{t+1+j} + (E_{t+1}-E_t)\sum_{j=0}^{\infty}\rho^j\Delta y_{t+1+j} + e_{t+1}. \quad (17)$$

We also assume that expected returns on the base assets are linear functions of Z_t, a vector of control variables known at time t:

$$E_t(R_{t,t+1}) = bZ_t. \quad (18)$$

And, we define the projection equation of lagged expectations of long-run labor income growth on the lagged control variables as

$$E_t\sum_{j=0}^{\infty}\rho^j\Delta y_{t+1+j} = fZ_t + \zeta_t. \quad (19)$$

Combining Eqs. (16)–(19) yields the following representation:

$$\sum_{j=0}^{\infty}\rho^j\Delta y_{t+1+j} = cR_{t,t+1} + dZ_t + \varepsilon_{t+1}, \quad (20)$$

where c and d are the regression coefficient vectors to be estimated, $\varepsilon_{t+1} \equiv \eta_{t+1} + e_{t+1} + \zeta_t$, ΔY_{t+1+j} is the labor income growth between $t + j$ and $t + 1 + j$, and $\rho = 0.954^{1/4}$.[6] Since the terms beyond a certain lead, S, on the left-hand side of Eq. (20) can be ignored,[7] Eq. (20) can be approximately rewritten as

$$\sum_{j=0}^{S}\rho^j\Delta y_{t+1+j} \cong cR_{t,t+1} + dZ_t + \varepsilon_{t+1}. \quad (21)$$

In fact, we consider several possible upper limits in the sum, and find no significant difference in results beyond 12 quarters. Thus, we set $S = 12$. We use quarterly data in estimating the regression model (21). Note that by including control variables on the right-hand side of the regression, we capture only the innovation component of future labor income. This is the critical difference between the economic tracking portfolio and the factor mimicking portfolio of Breeden et al. (1989).[8]

Returns on the tracking portfolio, which tracks innovations in future labor income growth, are computed by multiplying actual returns on the base assets by the regression coefficient, $\hat{c}$, estimated from Eq. (21). That is,

$$LIG_{t,t+1} = \hat{c}R_{t,t+1}. \quad (22)$$

According to the frequency of the base assets' returns, monthly or quarterly returns of the economic tracking portfolio are generated. The resulting economic tracking portfolio is the minimum variance combination of assets that is maximally correlated with future labor income growth. We use the *zero-investment* returns for the base assets, so that there is no restriction imposed on portfolio weights, c. Estimation of the tracking portfolios through Eq. (22) imposes no particular model of asset prices or equilibrium conditions. The only assumption used in deriving Eq. (21) is that information on changes in expectations about a future economic variable is reflected in asset returns, and these asset returns are a function of the lagged control variables. This assumption is justified if financial markets are efficient enough to reflect information on changes in expectations about future economic conditions.

Note that the use of the economic tracking portfolio is necessary for this study. The risk factor that our theoretical model implies is the *revision* in the expectation of future labor income growth, $(E_{t+1}-E_t)\sum_{j=0}^{\infty}\rho^j\Delta y_{t+1+j}$, rather than the expectation of or the actual future labor income growth. However, innovation in future labor income growth is unobservable. The construction of the economic tracking portfolio enables us to capture such unobservable components from asset returns, which are likely to contain information about the economic variable.[9]

3.2. Testing methodologies for the pricing ability of the risk factors

In order to examine whether revisions in the expectation of future labor income growth is priced in stock returns, we employ two estimation methods: the Fama and MacBeth (1973) two-pass methodology and the SDF approach implemented by the GMM estimation.

3.2.1. The Fama–MacBeth method

All test asset returns in excess of the riskless return are cross-sectionally regressed on their factor loadings estimates.[10] That is, for a given time t,

$$r_{i,t} = \lambda_{0t} + \lambda_t'\hat{\beta}_i + e_{i,t}, \quad (23)$$

where $r_{i,t}$ is the excess return on asset i; $\hat{\beta}_i$ is the $(K \times 1)$ factor loadings vector of asset i which are estimated in the first-pass intertemporal regression; $e_{i,t}$ is the error term; and λ_t is a $(K \times 1)$ parameter vector of the risk premia to be estimated at time t. The ultimate estimate of the risk premium for each risk factor is the time-series average of the month-by-month estimates of λ's, and its statistical significance is determined by the standard error of the time-series average. In the second pass cross-sectional regression (CSR), a well-known problem, the so called errors-in-variables (EIV) problem, arises due to the use of estimated factor loadings as regressors. To correct standard errors for the EIV problem, we use the Jagannathan and Wang's (1998) correction, since their approach is designed for the case of univariate regressions. Note that Shanken's (1992) correction is designed for the case of multiple regressions.[11]

To judge the overall fit of each asset pricing model in the CSR, we adopt the cross-sectional R^2 measure employed by Jagannathan and Wang (1996) and Lettau and Ludvigson (2001) as a summary statistic. This measure is defined as

$$R^2 = \frac{Var(\bar{r}) - Var(\bar{e})}{Var(\bar{r})}, \quad (24)$$

where $Var(\bar{r})$ is the cross-sectional variance of the average returns and $Var(\bar{e})$ is the cross-sectional variance of the residual average returns.

[6] Following the literature, we set ρ to be 5 % per annum, implying $\rho = 0.95^{1/4}$ quarterly. We also allow ρ to take different values between $0.9^{1/4}$ and 1, and find no significant difference in results.

[7] Due to the limited time-series observation of labor income series, we cannot continue the sum $\sum_{j=0}^{S}\rho^j\Delta y_{t+1+j}$ to the upper limit of infinity. However, it is likely that labor income growth very far out in the future will not matter substantially in estimating $\text{Cov}\left[r_{i,t+1}, (E_{t+1})\sum_{j=0}^{S}\rho^j\Delta y_{t+1+j}\right]$, because they are not likely to be correlated with current period returns.

[8] Another difference is that the dependent variable of Eq. (21) is the realized *future* economic variable, while the dependent variable in the estimation of the factor mimicking portfolio is the *contemporaneous* economic variable. Thus, the economic tracking portfolios are designed to capture information about future economic conditions.

[9] Alternatively, one can use vector autoregressive (VAR) approach to extract news about future labor income growth as in Campbell (1996). However, the economic tracing portfolio approach has a potential advantage over the VAR procedure. As the VAR system estimates factor loadings through a specific dynamic model of all the variables in the system, it can bring about a potential source of model misspecification. As Lamont (2001) argues, however, the tracking portfolio approach obtains data directly from the regression of the future economic variable on equity returns, without specifying a complete description of the data-generating process.

[10] We estimate the full-sample betas, as in Lettau and Ludvigson (2001) and Petkova (2006).

[11] Kim (1995, 1997) also suggests the EIV correction. Kim's correction is different from Shanken's (1992) and Jagannathan and Wang's (1998) in that his method corrects directly the CSR coefficient itself and compute the t-statistics from the corrected CSR coefficients, while their approaches correct only the t-statistics.

3.2.2. The stochastic discount factor approach

It is well known that when there is no arbitrage, there exists a positive stochastic discount factor (SDF) (or pricing kernel) m_{t+1} such that

$$E_t[m_{t+1}R_{t+1}] = 1_n, \tag{25}$$

where R_{t+1} is a $(n \times 1)$ vector of gross returns; 1_n is a $(n \times 1)$ vector of ones; and n is the number of test assets. Since all asset pricing models under consideration are linear factor pricing models, the pricing kernel can be represented as a linear combination of factors. That is,

$$m_{t+1} = b_0 + b_1' f_{t+1}, \tag{26}$$

where f_{t+1} is a $(K \times 1)$ vector of factors; b_0 is an intercept; and b_1 is a $(K \times 1)$ coefficient vector. b_0 and b_1 are called the SDF loadings.

In order to simultaneously estimate the tracking portfolios (i.e., estimating coefficients c and d in Eq. (21)) and the coefficients in the SDF (i.e., b_0 and b_1 Eq. (26)), the orthogonality condition of Eq. (21) is stacked at the moment condition of the asset pricing model of Eq. (25) such that

$$G(\theta) = \begin{pmatrix} E_t[\eta_{t,t+4} \otimes z_t] \\ E[m_{t+1}R_{t+1} - 1_n] \end{pmatrix} = \begin{pmatrix} 0 \\ 0 \end{pmatrix}, \tag{27}$$

where $\theta = (b_0, b_1, c, d)$ represents the parameters to be estimated, and $z_t = [B_{t-1,t}{:}\ Z_{t-2,t-1}]$ is a vector of the explanatory variables in Eq. (21). Since Eq. (27) is over-identified, the parameters estimates from the joint estimation can be different from the individual estimation. Following Vassalou (2003) and Aretz et al. (2010), we choose a matrix A so that both the joint and individual estimation produce the same parameter estimates. The matrix A is

$$A = \begin{bmatrix} I_{N_b+N_c} & 0 \\ 0 & d'\tilde{W} \end{bmatrix}, \tag{28}$$

where I is an identity matrix, N_b is the number of base assets, N_c is the number of control variables, d is the derivative matrix of the moment conditions of Eq. (25) with respect to the parameters in the SDF (i.e., b_0 and b_1), and $\tilde{E}$ is the weighting matrix used in two-step estimation. Parameters θ are chosen by minimizing the quadratic form

$$J_T = g(\theta)' W g(\theta), \tag{29}$$

where W is a weighting matrix. Since our competing models already have the determined risk factors, the estimation of the pricing ability of these competing models can be accomplished by Hansen's (1982) GMM method using Eq. (28).

Two weighting matrices are used to minimize the quadratic equation of Eq. (29). The first is the asymptotically optimal weighing matrix, which is adopted to compute Hansen's J-statistic on the overidentifying restrictions of the models. The second is the Hansen and Jagannathan (1997) weighing matrix, $E[RR']^{-1}$, which is the inverse of the second moments of asset returns; its main advantage is that it is invariant across competing asset pricing models. In order to compare the performance of pricing ability across models, therefore, we use this weighting matrix in computing the Hansen–Jagannathan distance (HJ-distance). The HJ-distance can be interpreted as the maximum pricing error for the set of assets mispriced by the model (Campbell and Cochrane, 2000).

According to Cochrane (1996), the risk premia, λ, can be estimated in the SDF approach as follows:

$$\lambda = -r_f \mathrm{Cov}(f, f') b_1, \tag{30}$$

where r_f is the riskless return; f is a $(K \times 1)$ vector of the factors; and b_1 is a $(K \times 1)$ coefficient vector in the pricing kernel of Eq. (26).

3.3. Data

For test assets, we use the 25 Fama and French size and book-to-market sorted portfolios, since these portfolios are one of the most commonly used test sets in the literature, due to their large cross-sectional dispersion in expected returns. These test assets and the three Fama–French factors are taken from Kenneth French's web site.[12] For labor income data, we use seasonally adjusted real per capita labor income from the second quarter of 1963 through the fourth quarter of 2007. This quarterly real per capital labor income data are taken from the National Income and Product Accounts (NIPA) Table 7.1, available from the Bureau of Economic Analysis. We make the standard "end-of-period" timing assumption that labor income during quarter t occurs at the end of the quarter.

The sample period of asset returns is accordingly determined by the availability of labor income data. Since we allow up 6 years in computing future labor income growth (i.e., $S = 16$ quarters in Eq. (21)), the actual test period is from the third quarter of 1963 through the fourth quarter of 2001.

Since there is little guidance on the choice of base assets to construct the economic tracking portfolio, it would be essential to follow a consistent rule in choosing base assets. Our principles are to choose a parsimonious set of base assets and a set of wide-range spanning the space of asset returns. These considerations lead us to choose ten industry portfolios as a first set of base assets.[13] Besides these equity portfolios, we include two bond market portfolios as a second set of base assets; DEF (the return difference between long-term corporate bonds and long-term government bond) and TERM (the return difference between long-term government bond and short-term government bond). These two bond portfolios are a widely-used choice in the related literature (e.g., Vassalou, 2003; Aretz et al., 2010, among others). To construct the risk factor, LIG, that captures revisions in the expectation of future labor income growth, therefore, we choose 10 industry portfolios plus TERM and DEF as base assets. Since we are interested in examining whether our proposed factor, LIG, can compete the Fama and French factors, SMB and HML, we do not include any portfolios containing information on the Fama and French benchmark factors. If the risk factor LIG is constructed by being nested with the Fama and French factors, the constructed factor could spuriously drive out the explanatory power of the Fama and French benchmark factors.

Construction of the tracking portfolios requires the control variables in order to capture unexpected components of base asset returns. Control variables should have the ability to predict future equity returns. Thus, we include the difference between the yields of a long-term corporate Baa bond and a long-term government bond (DEFY), the 3-month T-bill yield (RF), the difference between the yields of a 10-year and a one-year government bond (TERMY), and the consumption-wealth ratio (CAY) of Lettau and Ludvigson (2001). These variables are known for their predictive power.[14]

[12] We are grateful to Kenneth French for making the data available on his website:http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/data_library.html.

[13] Returns on these industry portfolios are obtained from Kenneth French's web site.

[14] These conditioning variables are also incorporated in estimating asset pricing models. Abhyankar and Gonzalez (2009) use the interest rate and the Baa–Aaa credit spread in testing their bond ICAPM. Moerman and van Dijk (2010) test a conditional version of the International Capital Asset Pricing Model relying on the default premium (the yield differential between Moody's Baa and Aaa rated bonds) and the term premium (the yield differential between the 10-year Treasury note and the Federal Funds Rate). Finally, Viale et al. (2009) employ term spread, default spread, and one-month Treasury bill yield in examining whether ICAPM explains bank stock returns.

All bond yield data are from the FRED® database of the Federal Reserve Bank of St. Louis.

4. Empirical results

4.1. Predictability of the base assets for future labor income growth

One necessary condition for selecting base assets is that base assets should reflect revisions of future expected income growth. It is important, therefore, to examine whether the chosen base assets are actually able to predict future labor income growth. In order to do so, we regress discounted future labor income growth rates from quarter $t+1$ to $t+1+S$ (i.e., $\sum_{j=0}^{S}\rho^j\Delta y_{t+1+j}$ in Eq. (21)) on contemporaneous returns of the four base assets and the control variables as in Eq. (21).

In order to determine a reasonable value for S that provides reliable and stable estimated value of the regression coefficients on the base asset returns, we estimate the regression model of Eq. (21) for $S=1$, 2, 4, 8, 12, 16, 20, and 24 quarters. Table 1 reports the estimation results of the regression coefficients for each value of S. Based on the p-values of the χ^2 tests for the significance of the estimated regression coefficients on the base assets, we choose $S=12$. Therefore, we use 12-quarter discounted future labor income growth rates to estimate returns on the tracking portfolio. For the robustness check, we have tested the longer horizons such as $S=16$, 20, and 24 quarters. However, the overall test results are qualitatively the same. The results with $S=16$ quarters are reported and discussed in Section 5 as a robustness check. Since the adjusted R^2 can be easily increased simply by including more control variables, the adjusted R^2 alone may not be a sufficient indicator of the tracking ability of the base assets. In order to fur-

Table 1

Predictability of future labor income growth by the base assets. The table reports the forecasting regression results for the following regression specification:

$$\sum_{j=0}^{S}\rho^j\Delta y_{t+1+j} = a + cB_{t-1,t} + dz_{t-2,t-1} + \eta_{t,t+s},$$

where $\sum_{j=0}^{S}\rho^j\Delta y_{t+1+j}$ is the discounted sum of labor income growth rate over the following S quarters, $B_{t-1,t}$ is the quarterly excess returns (over the T-bill rate) on the base assets (10 industry portfolios, the return difference between long-term corporate bonds and long-term government bond (DEF), the return difference between long-term government bond and short-term government bond (TERM)). $Z_{t-2,t-1}$ is the lagged control variables containing the difference between the yields of a long-term corporate Baa bond and a long-term government bond (DEFY), the 3-month T-bill yield (RF), the difference between the yields of a 10-year and a one-year government bond (TERMY), and the consumption-wealth ratio (CAY). The t-statistics are reported in parentheses and are corrected for White's (1980) serial correlation and heteroskedasticity using the Newey and West (1987a) estimator. The χ^2 statistics and their corresponding p-values are reported to test the null hypothesis that the coefficients of the base assets are jointly zero. The lower bond R^2 is obtained by regressing of changes in labor income growth expectations on unexpected stock returns. The sample period is from 1963:Q3 to 2007:Q4.

S	1	2	4	8	12	16	20	24
Base assets								
NoDur	−0.01	0.05	0.01	0.01	0.13	−0.06	−0.07	−0.03
	−0.11	0.90	0.12	0.09	1.01	−0.42	−0.43	−0.17
Durbl	0.00	0.04	0.04	0.10	0.12	0.12	0.08	0.10
	−0.16	1.13	0.92	1.56	1.61	1.47	0.96	1.21
Manuf	0.00	0.00	0.02	−0.07	−0.15	−0.12	−0.19	−0.21
	0.02	0.07	0.29	−0.72	−1.23	−0.97	−1.49	−1.48
Enrgy	−0.02	−0.02	−0.03	−0.02	−0.04	−0.05	−0.03	−0.03
	−2.46	−1.49	−1.30	−0.72	−1.37	−1.84	−0.97	−0.99
HiTec	0.01	−0.02	−0.06	−0.07	−0.06	−0.05	−0.03	−0.01
	0.79	−0.99	−2.85	−2.30	−1.74	−1.30	−0.72	−0.12
Telcm	0.01	0.03	0.05	0.06	0.06	0.02	0.04	0.02
	1.14	1.68	2.64	1.73	1.79	0.59	0.94	0.37
Shops	−0.01	−0.07	−0.05	0.02	−0.10	0.00	0.03	−0.01
	−0.19	−1.70	−0.83	0.19	−1.00	0.00	0.26	−0.13
Hlth	0.00	0.01	0.03	0.01	0.00	0.03	−0.02	−0.02
	−0.09	0.66	1.60	0.42	−0.10	0.73	−0.53	−0.63
Utils	0.05	0.03	0.00	−0.02	0.01	0.03	−0.01	−0.02
	2.26	1.02	−0.02	−0.42	0.14	0.47	−0.14	−0.25
Other	0.00	−0.01	0.00	−0.02	0.06	0.13	0.22	0.23
	0.15	−0.40	0.07	−0.26	0.82	1.47	2.50	2.45
DEF	0.00	0.09	0.11	0.02	−0.11	−0.09	−0.05	0.01
	0.03	1.63	1.43	0.11	−0.78	−0.58	−0.35	0.08
TERM	−0.04	0.00	−0.02	0.05	−0.02	−0.07	−0.07	−0.05
	−1.63	0.00	−0.43	0.88	−0.29	−0.99	−0.90	−0.58
Control variables								
Constant	1.06	2.09	3.47	5.92	8.19	10.25	12.60	15.20
	3.77	4.97	5.40	6.09	7.14	8.40	9.78	11.05
DEFY	−0.03	−0.04	0.14	0.59	0.68	−0.23	−0.57	−0.77
	−0.18	−0.19	0.43	1.23	1.32	−0.39	−0.89	−1.21
RF	−0.08	−0.14	−0.24	−0.37	−0.39	−0.20	−0.21	−0.30
	−1.52	−2.17	−2.78	−2.88	−2.83	−1.40	−1.40	−1.92
TERMY	0.05	−0.01	0.03	−0.28	−0.90	−0.85	−1.00	−1.15
	0.39	−0.07	0.16	−1.02	−2.99	−2.64	−2.90	−3.16
CAY	−0.10	−0.13	−0.23	−0.31	−0.18	−0.37	−0.39	−0.37
	−1.73	−1.71	−1.80	−1.76	−0.91	−1.52	−1.64	−1.51
R^2	0.16	0.15	0.20	0.19	0.18	0.16	0.21	0.23
Lower bound R^2	0.11	0.09	0.09	0.09	0.10	0.08	0.09	0.10
χ^2 (10)	18.01	13.74	16.19	17.56	23.27	16.20	22.67	24.66
p-value	0.12	0.32	0.18	0.13	0.03	0.18	0.03	0.02

Table 2
Descriptive statistics. The table reports the mean, standard deviation, and first-order autocorrelation of log consumption growth rate (CONS), excess market return (MKT), factor reflecting revisions in the expectation of future labor income growth (LIG), Fama–French factors related to size (SMB) and book-to-market (HML). It also reports the correlation among these variables. All data are quarterly observations. If portfolios' returns are given as monthly, these are compounded to convert into quarterly returns. The sample period is from 1963:Q3 to 2007:Q4.

Factors	CONS	MKT	LIG	SMB	HML
	Mean				
	0.57	1.34	0.02	0.69	1.33
	Standard deviation				
	0.45	8.57	1.02	5.99	6.03
	Autocorrelation				
	0.42	0.02	0.29	−0.02	0.14
	Correlation coefficient				
CONS		0.17	0.03	0.12	−0.03
MKT			0.23	0.49	−0.49
LIG				0.18	0.18
SMB					−0.19
HML					

ther examine the tracking ability of the base assets, therefore, we also compute the lower-bound adjusted R^2.[15] The lower bound R^2's are significantly large relatively to other studies, indicating that our tracking portfolio's returns do track relatively well innovations in the expectation of future labor income growth.

After estimating the regression coefficients of Eq. (21), returns on the tracking portfolio are obtained by multiplying the estimated regression coefficients by returns on the base assets, as in Eq. (22). We regard these returns as a risk factor associated with revisions in the expectation of future labor income growth and denote it as LIG. Table 2 reports the summary statistics of the five risk factors considered: excess market returns (MKT), log consumption growth rate (CONS), LIG, and Fama and French's SMB and HML. The average of LIG is insignificantly positive. More importantly, however, LIG is significantly correlated with the Fama–French factors: the correlation coefficients of LIG with SMB and HML are 0.18 and 0.18, respectively. This indicates that our proposed risk factor, LIG, shares information with the Fama–French factors.

4.2. Asset pricing test results

4.2.1. The pattern of the factor loadings on the tracking portfolios

It is widely accepted in the literature that firm size and book-to-market ratio are important forces in explaining stock returns. If a given factor is a determinant of average returns, then the loading associated with that factor should have a systematic pattern across firm sizes and book-to-market ratios. In this context, we examine whether there is a systematic pattern in the loading with the factor associated with revisions in the expectation about future labor income growth (LIG) across firm sizes and book-to-market ratios. Note that cross-sectional regression tests in this section are performed using quarterly returns since quarterly consumption growth rates are used.

Table 3 shows the estimation results of the time-series univariate regression model of each of the 25 Fama and French size/BM-sorted portfolios on each of the three factors. The factor loadings associated with revisions in the expectation of future labor income growth (β_{LIG}) are significantly estimated. More importantly, the estimated factor loadings show a systematic pattern across both firm size and book-to-market. That is, β_{LIG} monotonically decreases with firm size within each book-to-market quintile and increases with book-to-market within each firm size quintile. Thus, this is indirect evidence that the future labor income growth factor is related to both firm size and book-to-market. This is an interesting result, since we find that each of the Fama and French factors is related only to its own corresponding characteristic, that is, SMB is related only to firm size, and HML is related only to book-to-market.[16]

One necessary condition for a factor loading to have satisfactory explanatory power for cross-sectional variations in average returns is for it to have sufficient cross-sectional spread in factor loading. In this sense, the future labor income growth factor satisfies this necessary condition, since the magnitude of the cross-sectional spread in β_{LIG} is greater than that in the other factor loadings under consideration. For example, the cross-sectional spread in β_{LIG} is between 1.27 and 4.31. However, the cross-sectional spreads in β_{SMB} and β_{HML} are only between −0.28 and 1.59 and between −0.42 and 1.02, respectively (unreported).[17]

4.2.2. Results of cross-sectional regression tests

In the time-series tests, we preliminarily observe a positive cross-sectional association between the factor loadings on the future labor income growth factor (β_{LIG}) and average returns. In order to formally examine whether the risk associated with revisions in the expectation of future labor income growth is priced, we perform cross-sectional regression tests.

Table 4 reports the CSR estimation results of the Fama and French three-factor model (in Panel A), our alternative three-factor model (in Panel B), a one-factor model including LIG only (in Panel C), a five-factor model including all five factors (in Panel D), and a three-factor model including LIG, SMB, and HML (in Panel E), within Fama and MacBeth's (1973) methodology framework. Test portfolios are the 25 Fama and French size and book-to-market sorted portfolios, and betas are estimated in the first-pass univariate time-series regression models over the whole test period from 1963:Q3 to 2001:Q4.[18]

Panel A of Table 4 shows that the Fama and French three-factor model shows a significant explanatory power. The adjusted R^2 is 0.74 and the risk premium estimate of HML is positively significant, while that of SMB is insignificant. Panel B shows that our alternative model also has a significant explanatory power. The consumption growth beta and the market beta are not significant, but the future labor income growth beta is statistically positively significant. The risk premium of LIG is estimated as 0.85% (with t-statistic of 3.84). Moreover, the intercept estimate of our alternative model is not statistically significant. The intercept estimate is 0.52%, with t-statistic of 0.56. Our alternative model performs slightly better than the Fama and French three-factor model in terms of adjusted R^2. The adjusted R^2 of our alternative model is 0.81. This is somewhat surprising when we recall that SMB and HML are constructed with the same characteristics as the test portfolios. When the revisions in the expectation of future labor in-

[15] The lower-bound R^2 is a partial R^2 that gives a lower bound on the fraction of the variance of innovations that is captured by tracking portfolio returns. It is calculated as follows. We first regress the future labor income growth rates (the Y-variable in the regression equation (19)) onto our control variables. Subsequently, we regress the tracking portfolio return (LIG) onto our control variables. Finally, we regress the residuals from the former regression onto the residuals from the latter regression. The R^2 in this regression is lower bound for the R^2 of the full regression of Eq. (19).

[16] More specifically, as designed, the factor loadings on SMB (β_{SMB}) show a monotonic decreasing pattern across firm size but almost no pattern across book-to-market. Meanwhile, the factor loadings on HML (β_{HML}) show a monotonic increasing pattern across book-to-market but almost no pattern across firm size.

[17] The results for β_{SMB} and β_{HML} are available upon request.

[18] It is worth noting that since the covariances or betas in the derivation of our three-factor model are separately considered; the use of univariate betas in asset pricing tests is consistent with the theoretical derivation. It is important, however, to note that our results remain quantitatively the same even when betas are jointly estimated in multivariate regressions.

Table 3
Factor loading estimates of the risk factors. This table reports the estimate results of the time-series univariate regression of returns on the Fama–French 25 size and book-to-market sorted portfolios on each of the three risk factors: excess market returns (MKT), log consumption growth rate (CONS), revisions in the expectation of future labor income growth (LIG). *t*-statistics are corrected for autocorrelation and heteroskedasticity using the Newey and West (1987a) estimator with three lags. The sample period is from 1963:Q3 to 2007:Q4.

	Low	2	3	4	High	Low	2	3	4	High
	$\beta_{\Delta c}$(CONS)					$t(\beta_{\Delta c})$				
Small	6.72	6.52	4.83	4.90	5.44	2.24	2.55	2.09	2.25	2.30
2	4.75	3.94	4.14	3.63	4.60	1.76	1.78	2.14	1.89	2.27
3	3.85	3.63	3.56	3.22	3.79	1.52	1.78	1.99	1.66	2.09
4	3.67	3.45	2.60	2.93	3.80	1.52	1.84	1.44	1.71	1.87
Large	3.27	1.93	2.98	2.07	3.02	1.77	1.23	2.10	1.41	2.24
	β_M(MKT)					$t(\beta_M)$				
Small	1.66	1.40	1.23	1.15	1.19	19.86	19.21	16.33	15.63	14.56
2	1.57	1.29	1.15	1.05	1.08	21.51	19.46	17.30	15.54	13.06
3	1.47	1.18	1.01	0.95	0.98	23.86	23.43	15.13	14.15	11.49
4	1.34	1.10	0.96	0.93	0.98	23.54	18.72	17.56	16.71	12.66
Large	1.03	0.91	0.76	0.75	0.75	40.42	23.36	17.76	15.28	13.63
	β_{LIG}(LIG)					$t(\beta_{LIG})$				
Small	2.44	3.25	3.50	3.54	4.31	1.51	2.45	3.11	3.43	3.79
2	2.34	3.27	3.09	3.48	3.98	1.72	2.88	3.19	3.76	4.24
3	2.42	2.83	2.93	2.99	3.63	2.02	2.90	3.57	3.46	4.12
4	1.64	2.29	2.52	2.89	3.50	1.49	2.45	3.13	3.70	4.22
Large	1.75	1.79	1.27	2.45	2.82	2.15	2.23	1.65	3.85	4.16

come risk factor alone is included in the model (in Panel C), the adjusted R^2 is 0.78. This indicates that LIG contributes most of the explanatory power of our alternative three-factor model.

In order to examine whether there is an incremental explanatory power for SMB and HML when they are added into our three-factor model, we estimate a five-factor model including all five factor loadings (in Panel D) as well as a three-factor model including SMB, HML, and LIG only (in Panel E). The CSR results show that only the coefficient estimate on β_{LIG} is significant and has an economically consistent sign; it is 0.77% (with *t*-statistic of 3.12) in the five-factor model. However, the coefficient estimates on β_{SMB} and β_{HML} are all insignificant. Moreover, the statistical significance of HML disappears, when LIG is included in the model. These results are also confirmed when we estimate a three-factor model including SMB, HML, and LIG (in Panel E). These results imply that SMB and HML have no incremental explanatory power for the cross-section of average returns when the future labor income growth factor is included in the model. It could be argued, therefore, that revisions in the expectation of future labor income growth absorb the pricing effect of SMB and HML.[19] One noteworthy consideration is that the magnitude of the risk premium of LIG is very stable across the estimated models. Since the CSR coefficient estimates are subject to the errors-in-variables (EIV) problem, we also report the EIV-corrected *t*-statistics by using Jagannathan and Wang's (1998) approach. However, the results are not qualitatively changed.

4.2.3. GMM estimation results

Along with the Fama–MacBeth CSR tests, we evaluate the performance of our alternative models using the SDF approach implemented through the GMM estimation. Table 5 reports the GMM estimation results using the optimal weighting matrix, which are generally consistent with the CSR results. The GMM estimation results for the Fama and French three-factor model (in Panel A) show that SMB and HML are significantly priced. Their SDF loadings and risk premia are statistically significant, and the *p*-values of the

[19] As a robustness test, we run the CRS with beta estimates of orthogonalized factors to examine whether the corresponding factor is marginally useful in pricing assets, given the presence of other factors. We find that only the coefficient of the orthogonalized factor associated with labor income risk is significant.

Table 4
Cross-sectional regression estimation results. The table reports the time-series averages (in percent per quarter) of the regression coefficient estimates of the cross-sectional regression model:

$$r_{i,t} = \lambda_{0t} + \lambda_t'\hat{\beta}_i + e_{i,t},$$

where $r_{i,t}$ is the return of portfolio *i* in excess of the riskless return, and $\hat{\beta}$ is the factor loadings estimated in the first-pass univariate time-series regression model using quarterly returns over the whole sample period. The test portfolios are Fama–French's (1993) 25 portfolios independently sorted by size and book-to-market. CONS is the log consumption growth rate, MKT is the market return in excess of the riskless rate of return, SMB and HML are Fama and French's (1993) factors related to firm size and book-to-market, and LIG is the factor reflecting revisions in the expectation of future labor income growth. "*t*-value" is computed by using the uncorrected Fama–MacBeth standard errors. "JW corrected-*t*" is computed by using Jagannathan and Wang's (1998) correction for the errors-in-variables bias. The adjusted R^2 is computed by using Jagannathan and Wang's (1996). The sample period is from 1963:Q3 to 2007:Q4.

	Constant	MKT	SMB	HML			Adj. R^2
Panel A: The Fama–French three-factor model							
Estimate	1.05	1.17	0.64	1.87			0.74
t-value	0.69	0.43	0.66	2.03			
JW corrected-*t*	0.69	0.45	0.68	2.09			
		CONS	MKT	LIG			
Panel B: Our alternative three-factor model							
Estimate	0.52	0.04	−0.82	0.85			0.81
t-value	0.56	0.29	−0.75	3.84			
JW corrected-*t*	0.39	0.23	−0.45	2.84			
		LIG					
Panel C: A one-factor model							
Estimate	−0.26	0.86					0.78
t-value	−0.30	3.45					
JW corrected-*t*	−0.21	3.06					
		CONS	MKT	LIG	SMB	HML	
Panel D: A five-factor model							
Estimate	1.64	−0.03	−2.24	0.77	0.66	−0.26	0.80
t-value	1.14	−0.20	−0.83	3.12	0.73	−0.25	
JW corrected-*t*	0.87	−0.15	−0.61	2.32	0.66	−0.17	
		LIG	SMB	HML			
Panel E: A three-factor model							
Estimate	0.38	0.69	0.08	0.49			0.81
t-value	0.54	3.02	0.13	0.82			
JW corrected-*t*	0.44	2.46	0.12	0.64			

Table 5
GMM estimation results. This table reports the GMM estimation results by using Fama and French's (1993) 25 size and book-to-market sorted portfolios. CONS is the log consumption growth rate, MKT is the market return in excess of the riskless rate of return, SMB and HML are Fama and French's (1993) factors related to firm size and book-to-market, and LIG is the factor reflecting revisions in the expectation of future labor income growth. All data are quarterly observations. If portfolios' returns are given as monthly, these are compounded to convert into quarterly returns. *t*-values in parentheses are computed from the one-step estimation, where the tracking portfolio and the asset pricing model are simultaneously estimated. *t*-values in square brackets are computed from the two-step estimation. The Wald (*b*) test is a joint significance test of the factor loadings in the pricing kernel. The SDF loadings and the test statistics for Wald (*b*) are computed through the GMM estimation that uses the optimal weighting matrix. The HJ-distance is the Hansen–Jagannathan (1997) distance measure, and its *p*-value is obtained from 10,000 simulations. The sample period is from 1963:Q3 to 2007:Q4.

	Constant	MKT	SMB	HML				Wald (*b*)
Panel A: The Fama–French three-factor model								
Factor loadings	1.10	−0.01	−0.03	−0.05			Test statistic	18.77
[*t*-value]	21.91	−0.60	−1.49	−2.83			*p*-value	0.00
								HJ-distance
Risk premium		0.48	1.01	1.15			Test statistic	0.63
[*t*-value]		0.50	2.49	2.63			*p*-value	0.00
		CONS	MKT	LIG				Wald (*b*)
Panel B: The alternative three-factor model								
Factor loadings	0.84	0.20	0.02	−0.83			Test statistic	33.87
[*t*-value]	4.11	0.56	1.36	−5.69			*p*-value	0.00
(*t*-value)	1.95	0.27	0.54	−2.39				
								HJ-distance
Risk premium		−0.04	0.36	0.86			Test statistic	0.60
[*t*-value]		−0.54	0.40	5.62			*p*-value	0.00
(*t*-value)		−0.24	0.15	2.37				
		CONS	MKT	LIG	SMB	HML		Wald (*b*)
Panel C: A five-factor model								
Factor loadings	1.03	−0.08	0.02	−0.65	0.00	−0.02	Test statistic	33.04
[*t*-value]	4.39	−0.20	0.60	−3.04	−0.09	−0.99	*p*-value	0.00
(*t*-value)	3.51	−0.15	0.44	−1.76	−0.09	−0.62		
								HJ-distance
Risk premium		0.02	−0.22	0.70	0.32	1.91	Test statistic	0.58
[*t*-value]		0.19	−0.19	3.53	0.66	4.55	*p*-value	0.00
(*t*-value)		0.15	−0.15	2.10	0.53	2.69		

Wald (*b*) tests are less than a 1% significance level.[20] Note that the Wald (*b*) test examines whether the coefficients in the pricing kernel, *b*, or the SDF loadings are jointly zero. Rejecting the null hypothesis of $b = 0$ implies that the factors jointly have important implication of the SDF and have marginal explanatory power for pricing the test portfolios.

For our three-factor model (in Panel B), the SDF loading on LIG is statistically significant (with *t*-statistic of −5.69), implying that revisions in the expectation of future labor income growth are important in explaining stock returns. The risk premium on LIG (λ_{LIG}) is statistically significant; it is 0.86% with *t*-statistic of 5.62. The *p*-value of the Wald (*b*) test for our three-factor model is less than 0.01, rejecting the null hypothesis of $b = 0$. Panel C shows the GMM estimation results of the five-factor model. When LIG is included in the model, the significance of the SDF loadings on SMB and HML disappears, and the SDF loading on LIG only is significant. Moreover, the risk premium on LIG is still significant. In the five-factor model, the risk premium on HML is also significant. However, when the identity matrix or Hansen and Jagannathan (1997) weighting matrix is used in the GMM estimation instead of the optimal weighting matrix, the risk premium on HML is not significant, while the risk premium on LIG is still significant. These results are consistent with the CSR results. The *t*-values of the coefficient estimates on the models including LIG are computed both from the one-step joint estimation where the tracking portfolio and the asset pricing model are simultaneously estimated and from the two-step estimation.

[20] The loading on SMB is moderately significant, although it is not significant at the traditional significance level such as one or five percent level. For this pattern see also Kim (2010).

To compare the performance of the asset pricing models, we compute the HJ-distance, which translates into the maximum pricing error generated by each of the models. In terms of the HJ-distance, our model performs better in explaining the cross-section of returns than the Fama and French model. The values of the HJ-distance are 0.60 and 0.63 for our alternative model and the Fama–French model, respectively. These results are consistent with the CSR tests in which the adjusted R^2 of our alternative model is higher than the Fama–French model. Nonetheless, the HJ-distance test for the null hypothesis that the squared pricing errors are statistically different from zero rejects both our model (*p*-value = 0.00) and the Fama–French model (*p*-value = 0.00) at a 5% significance level, implying that any of the models considered does not correctly price the test assets.

4.3. Relations between the Fama–French factors and the future labor income growth factor

The CSR testing results shown above show that our alternative model performs better than the Fama and French three-factor model, and that our proposed factor, the future labor income growth factor, subsumes the explanatory power of the Fama–French factors in explaining the cross-section of stock returns. In order to more directly examine whether the future labor income growth factor, LIG, shares important pricing information with the Fama and French factors, we run the following time-series regression equations:

$$\text{LIG}_t = \alpha_{\text{LIG}} + \delta^{\text{S}}_{\text{LIG}}\text{SMB}_t + \delta^{\text{H}}_{\text{LIG}}\text{HML}_t + \varepsilon_t, \tag{31}$$

$$\text{SMB}_t = \alpha_{\text{SMB}} + \delta_{\text{SMB}}\text{LIG}_t + \varepsilon_t, \tag{32}$$

$$\text{HML}_t = \alpha_{\text{HML}} + \delta_{\text{HML}}\text{LIG}_t + \varepsilon_t. \tag{33}$$

Table 6
Pricing the Fama–French Factors. The table reports the estimation results of time-series regressions of factor *j* on the other factors by using monthly or quarterly data. SMB and HML are Fama and French's (1993) factors related to firm size and book-to-market, and LIG is the factor reflecting revisions in the expectation of future labor income growth. *t*-statistics are reported below the coefficient estimates and are corrected for autocorrelation and heteroskedasticity using the Newey–West (1987b) estimator with three lags. All data are quarterly observations. If portfolios' returns are given as monthly, these are compounded to convert into quarterly returns. The sample period is from 1963:Q3 to 2007:Q4.

	Constant	SMB	HML	Adj. R^2
Panel A: $LIG_t = \alpha_{LIG} + \delta^S_{LIG} SMB_t + \delta^H_{LIG} HML_t + \varepsilon_t$				
Quarterly				
Estimate	−0.05	0.04	0.04	0.07
t-value	−0.48	2.44	2.30	
Monthly				
Estimate	−0.01	0.02	0.04	0.05
t-value	−0.48	1.77	3.47	
	Constant	LIG		
Panel B: $SMB_t = \alpha_{SMB} + \delta_{SMB} LIG_t + \varepsilon_t$				
Quarterly				
Estimate	0.66	1.08		0.03
t-value	1.30	1.90		
Monthly				
Estimate	0.22	0.31		0.00
t-value	1.38	0.65		
	Constant	LIG		
Panel C: $HML_t = \alpha_{HML} + \delta_{HML} LIG_t + \varepsilon_t$				
Quarterly				
Estimate	1.31	1.07		0.03
t-value	2.45	1.79		
Monthly				
Estimate	0.42	1.28		0.04
t-value	2.69	2.95		

Table 6 reports the coefficient estimates of the above time-series regression models along with the corresponding *t*-statistics corrected for heteroskedasticity and autocorrelation. The coefficient estimates on SMB and HML in Eq. (31), δ^S_{LIG} and δ^H_{LIG} are all positive and statistically significant for both monthly and quarterly data. The coefficients on LIG in Eqs. (32) and (33), δ_{SMB}, and δ_{HML}, are also positive and statistically significant for both monthly and quarterly data, except for δ_{SMB} for monthly data. Over all, there is a significantly positive association between LIG and the Fama and French factors. These results indicate that the risk factor reflecting revisions in the expectation of future labor income growth covaries positively with SMB and HML.

In order to more thoroughly examine the relationship between the Fama and French factors and the future labor income growth factor, we investigate the time-series movements of the risk premiums of these factors with respect to business cycles. The risk premium associated with the combined Fama and French factors is obtained as follows: First, each of the 25 Fama and French portfolios is regressed on MKT, SMB, and HML. Second, we compute the average of the 25 Fama and French portfolios' estimated corresponding factor loadings, multiplied by SMB and HML (that is, $\hat{\beta}_{i,SMB}SMB_t + \hat{\beta}_{i,HML}HML_t$); this is regarded as the combined risk premium for the Fama and French factors. Fig. 1 plots the time-series movement of these two risk premia, and shows a quite positive association between the risk premia. In particular, in bad state of the economy, both risk premia are higher, and they covary more closely than in good state of the economy. Specifically, the correlation coefficients over the whole, contraction, and expansion periods are 0.26, 0.55, and 0.16, respectively. It is interesting that the correlation between the two risk premia is higher in a contracting period than in an expanding period. Note that we use the NBER definition of the business cycle, and the shaded bar in Fig. 1 indicates the contraction period.[21]

In fact, a positive relationship between LIG and the Fama–French factors is economically plausible. When good states of the economy are expected, small capitalization stocks and value stocks with high financial leverage might be able to better prosper than big capitalization stocks and growth stocks. As a result, during good (bad) times in terms of business conditions, when the future labor income growth rate is expected to be high (low), returns on small stocks and high book-to-market stocks would be relatively higher (lower) than those of big stocks and low book-to-market stocks. Thus, returns on SMB and HML are positively associated with shocks to the level of human capital. We interpret these results as suggesting that small firms and value stocks are more sensitive to shocks to the state of the human capital. That is, small stocks and value stocks are indeed fundamentally riskier than big stocks and growth stocks.

Another possible interpretation for the positive association between news about future labor income growth and the Fama–French factors is asymmetric employment across firms. In recessions, employment in small and value firms, typically weak firms with persistently low earnings and high cash flow uncertainty (Chan and Chen, 1991; Fama and French, 1995), is more likely to contract than in big and growth firms. Thus, a negative shock to SMB and HML more likely implies a negative shock to the value of human capital. Rational investors, who have hedging concerns with respect to their future labor income, have an incentive to avoid the stocks of small and value firms. As a result, small and value firms are riskier than big and growth firms in recessions, when the price of risk associated with labor income is high. Indeed, our results echo the view of Fama and French (1996, p. 77): "Why is relative distress a state variable of special hedging concern to investors? One possible explanation is linked to human capital, an important asset for most investors."

4.4. Comparison with competing asset pricing models

In order to examine how well our three-factor model performs in explaining the cross-section of stock returns, we estimate several competing models, using the CSR and GMM estimations. Table 7 reports the CSR and GMM estimation results of six competing models: the CAPM (in Panel B), the Jagannathan and Wang (1996) human capital CAPM (in Panel C), the consumption CAPM of Breeden (1979) (in Panel D), the Epstein–Zin (1991) model (in Panel E), and the Lettau–Ludvigson (2001) model (in Panel F).The results of our three-factor model from Tables 4 and 5 are repeated in Panel A. The human capital CAPM is regarded as comparing the pricing ability of the factors related to *current* labor income growth and to innovations to *future* labor income growth. The Lettau–Ludvigson model is considered because it uses macroeconomic variables similar to ours and is known to have an almost equal ability in explaining the cross-section of average stocks returns as the Fama and French three-factor model.

The overall results of Table 7 show that our three-factor model performs better than any other competing model. The HJ-distance of our three-factor model is 0.60, which is the smallest among those of all competing models. Note that the HJ-distance measures (maximum) pricing errors. The values of the HJ-distance are 0.68,

[21] We also obtain similar results when we use a different definition of business conditions. As in Petkova and Zhang (2005), we define a time period as 'peak' if the expected risk premium of the period is below the bottom 10%, and as 'trough' if the expected risk premium of the period is above the top 10% among the whole periods. The expected risk premium is obtained as the fitted value of the *Y*-variable from the regression of the market excess return on the lagged TERM, DEF, and the risk-free return.

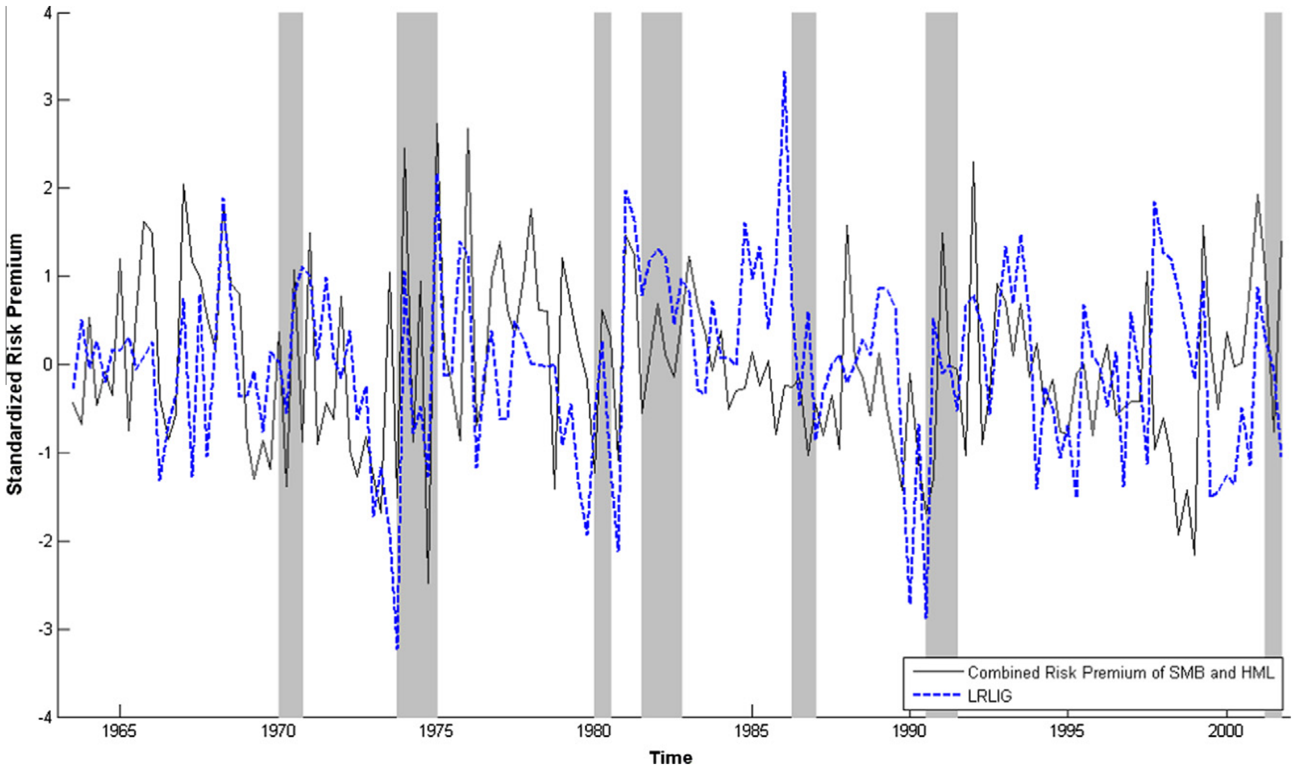

Fig. 1. Risk premium associated with the Fama–French Factors and the future labor income growth factor. The figure is a time-series plot of the risk premium associated with revisions in the expectation of future labor income growth (LIG) (dotted line) and the Fama–French factors (solid line). The risk premium associated with the combined Fama–French factors is obtained as follows: Each of the Fama and French 25 portfolios is regressed on MKT, SMB, and HML. Then, we compute the average of Fama and French 25 portfolios' estimated corresponding factor loadings times SMB and HML (that is, $\hat{\beta}_{i,\mathrm{SMB}}\mathrm{SMB}_t + \hat{\beta}_{i,\mathrm{HML}}\mathrm{HML}_t$), which is regarded as the combined risk premium for the Fama and French factors. Both series are normalized to standard deviations of unity. The shaded regions indicate the recession periods defined by NBER.

0.65, 0.68, 0.68, and 0.67, respectively, for the CAPM, the human capital CAPM, the consumption CAPM, the Epstein–Zin model, the Lettau–Ludvigson model. Recall that the HJ-distance of the Fama–French model is 0.63. The intercept estimate of our three-factor model is insignificant. However, the intercept estimates of the competing models are all quite significant.

The Wald (SMB&HML) statistic is used to test whether the Fama–French factors SMB and HML are marginally useful in pricing assets, given the presence of other factors, when these two factors are added into the model. The *p*-value of the Wald (SMB&HML) test for our three-factor model is 0.38, which means that when our three factors, especially LIG, are included in the model, the SDF loading estimates of SMB and HML turn out to be insignificant and their marginal explanatory power for the cross-section of stock returns is limited. However, the Wald (SMB&HML) test for the competing models indicates that there is still room for SMB and HML in explaining the cross-section of stock returns. Put differently, the factors of the competing models do not successfully explain the portion of the cross-section of average returns that SMB and HML do.

4.5. Comparison with innovations in future GDP growth

The previous sections show that revision in the expectation of future labor income growth, capturing changes in the value of human capital, is related to the Fama and French factors (SMB and HML). A recent paper by Vassalou (2003) also provides evidence that another important macroeconomic variable, future GDP growth is related to the Fama and French factors. Vassalou interprets her results as an Intertemporal Capital Asset Pricing (ICAPM) explanation of the Fama and French model. However, Petkova (2006) points out that it is unclear whether the GDP growth actually represents an ICAPM state variable, since Vassalou uses an *ad hoc* assumption that changes in the investment opportunity set can be described by changes in future GDP growth.

Nonetheless, it is worth to compare our future labor income growth with Vassalou's (2003) future GDP growth, since these two macroeconomic variables are related with each other. The procedure of constructing both factors is similar: both are constructed by regressing the macroeconomic variable of interest on returns of base assets. Therefore, a natural question arises as to whether revisions in the expectation of future labor income growth simply capture news about GDP growth. In order to examine this possibility, we do the CSR tests by regressing excess returns of the test portfolios on the factor loadings on LIG and GDPG (factor reflecting innovations in GDP growth rate).[22]

Panel A of Table 8 shows that GDPG is priced when both MKT and GDPG are in the model. However, this significance of GDPG disappears when LIG is added into the model, while the significance of LIG is still maintained (in Panel B). This indicates that LIG contains a substantial amount of explanatory power of GDPG for average stock returns. However, these results might come from a high correlation between labor income growth and GDP growth, since labor income is a part of GDP. In order to compare a marginal effect by these two macroeconomic variables, we construct the orthogonalized factors of LIG and GDPG as follows. We first regress

[22] More specifically, following Vassalou (2003), we regress GDP growth one-year ahead on the returns on base asset and lagged control variables. In order to fairly compare LIG and GDPG, we use the same base assets: ten industry portfolios plus TERM and DEF. When we use eight base assets (Fama and French's (1993) six portfolios and two bond portfolios), which are used in Vassalou (2003), our results also remain quantitatively similar.

Table 7
Comparison of competing asset pricing models. This table reports the CSR estimation results (in the upper part) and the GMM estimation results (in the lower part) of the competing models. CONS is the log consumption growth rate, MKT is the market return in excess of the riskless rate of return, SMB and HML are Fama and French's (1993) factors related to firm size and book-to-market, LIG is the factor reflecting revisions in the expectation of future labor income growth, LI is the log of (present) labor income growth rate, and CAY is the consumption-wealth ratio created by Lettau and Ludvigson (2001). All data are quarterly observations. If portfolios' returns are given as monthly, these are compounded to convert into quarterly returns. The risk premiums associated with factors are estimated using Fama–MacBeth method. "JW corrected-*t*" is computed by using Jagannathan and Wang's (1998) correction for the errors-in-variables bias. The adjusted R^2 is computed by using Jagannathan and Wang (1996). The Wald (*b*) test is a joint significance test of the factor loadings in the pricing kernel. The HJ-distance is the Hansen-Jagannathan (1997) distance measure, and its *p*-value is obtained from 10,000 simulations. The *J*-test is Hansen's (1982) test on the overidentifying restrictions of the model. The Wald (SMB&HML) statistic tests whether SMB and HML contain an incremental ability in pricing the test assets. The test statistics for Wald (*b*), HJ-distance, *J*-test, and Wald (SMB&HML) are computed through the GMM estimation. The sample period is from 1963:Q3 to 2007:Q4.

	Constant	CONS	MKT	LIG	Adj. R^2
Panel A: The alternative model					
Risk premium	0.52	0.04	−0.82	0.85	0.81
t-value	0.56	0.29	−0.75	3.84	
JW corrected-*t*	0.39	0.23	−0.45	2.84	
	Tests				
	Wald (*b*)	HJ-distance	*J*-test	Wald (SMB&HML)	
Statistic	33.87	0.60	63.39	1.95	
p-value	0.00	0.00	0.00	0.38	
	Constant	MKT			
Panel B: The CAPM					
Risk Premium	2.71	−0.48			−0.02
t-value	2.96	−0.42			
JW corrected-*t*	2.95	−0.42			
	Tests				
	Wald (*b*)	HJ-distance	*J*-test	Wald (SMB&HML)	
Statistic	8.93	0.68	89.95	7.43	
p-value	0.00	0.00	0.00	0.02	
	Constant	MKT	LI		
Panel C: The Jagannathan–Wang model					
Risk Premium	3.20	−1.14	−1.12		0.24
t-value	3.32	−0.99	−3.38		
JW corrected-*t*	2.08	−0.59	−2.63		
	Tests				
	Wald (*b*)	HJ-distance	*J*-test	Wald (SMB&HML)	
Statistic	13.48	0.65	59.46	5.02	
p-value	0.00	0.01	0.00	0.08	
	Constant	CONS			
Panel D: The consumption CAPM					
Risk premium	1.58	0.15			0.02
t-value	2.40	0.72			
JW corrected-*t*	2.16	0.68			
	Tests				
	Wald (*b*)	HJ-distance	*J*-test	Wald (SMB&HML)	
Statistic	0.05	0.68	115.86	8.22	
p-value	0.82	0.00	0.00	0.02	
	Constant	CONS	MKT		
Panel E: The Epstein–Zin model					
Risk premium	2.84	0.54	−2.49		0.29
t-value	3.10	2.65	−2.05		
JW corrected-*t*	2.40	1.42	−1.26		
	Tests				
	Wald (*b*)	HJ-distance	*J*-test	Wald (SMB&HML)	
Statistic	12.85	0.68	91.09	7.49	
p-value	0.00	0.00	0.00	0.02	
	Constant	CAY	CONS	CAY * CONS	
Panel F: The Lettau–Ludvigson model					
Risk premium	3.35	−0.02	0.05	0.01	0.54
t-value	3.17	−2.54	0.21	2.64	
JW corrected-*t*	2.91	−2.18	0.14	1.68	

Table 7 (*continued*)

	Constant	CONS	MKT	LIG	Adj. R^2
	Tests				
	Wald (*b*)	HJ-distance	*J*-test	Wald (SMB&HML)	
Statistic	11.10	0.67	98.16	7.37	
p-value	0.01	0.01	0.00	0.03	

Table 8
Comparison between GDP growth and labor income growth. The table reports the time-series averages of the cross-sectional regression coefficient estimates of the portfolio excess returns on the factor loadings. The test portfolios are Fama–French's (1993) 25 portfolios independently sorted by size and book-to-market. MKT is the market return in excess of the riskless rate of return, SMB and HML are Fama and French's (1993) factors related to firm size and book-to-market, GDPG is the factor reflecting news about future GDP growth, and LIG is the factor reflecting revisions in the expectation of future labor income growth. We construct the orthogonalized factors of LIG and GDPG as follows. We first regress GDP growth rates on labor income growth rates and take the intercept and the residuals. By regressing the intercept and the residuals on returns of the base assets, we then obtain the economic tracking portfolio. This is $GDPG^{\perp}$. Thus, $GDPG^{\perp}$ is the economic tracking portfolio capturing news about remaining GDP growth after excluding labor income component. Likewise, $LIG^{\perp}$ is LIG after excluding the effect by GDPG. "*t*-value" is computed by using the uncorrected Fama–MacBeth standard errors. "JW corrected-*t*" is computed by using Jagannathan and Wang's (1998) correction for the errors-in-variables bias. The adjusted R^2 is computed by using Jagannathan and Wang's (1996). The sample period is from 1963:Q3 to 2007:Q4.

	Constant	MKT	GDPG			Adj. R^2
Panel A: A two-factor model						
Estimate	−1.71	−0.90	0.79			0.42
t-value	−1.29	−0.79	3.48			
JW corrected-*t*	−0.87	−0.47	2.44			
		MKT	GDPG	LIG		
Panel B: A three-factor model						
Estimate	0.31	−0.68	0.04	0.85		0.81
t-value	0.30	−0.59	0.17	2.95		
JW corrected-*t*	0.21	−0.35	0.12	2.82		
		MKT	GDPG	SMB	HML	
Panel C: A four-factor model						
Estimate	1.77	−2.22	0.25	1.38	0.70	0.74
t-value	1.30	−0.83	1.24	1.61	0.71	
JW corrected-*t*	1.25	−0.81	1.19	1.50	0.72	
		MKT	GDPG	SMB	HML	
Panel D: A four-factor model including the orthogonalized GDPG						
Estimate	1.23	0.42	0.05	0.70	1.65	0.73
t-value	0.84	0.14	0.21	0.73	1.62	
JW corrected-*t*	0.84	0.14	0.21	0.75	1.64	
		MKT	LIG	SMB	HML	
Panel E: A four-factor model including the orthogonalized LIG						
Estimate	1.15	0.79	0.55	0.49	0.65	0.80
t-value	0.75	0.29	2.87	0.51	0.62	
JW corrected-*t*	0.58	0.22	2.28	0.46	0.46	

GDP growth rates on labor income growth rates and take the intercept and the residuals. By regressing the intercept and the residuals on returns of the base assets, we then obtain the economic tracking portfolio, as described in Section 3.1. We call this $GDPG^{\perp}$. Thus, $GDPG^{\perp}$ is the economic tracking portfolio capturing news about remaining GDP growth after excluding labor income component. Likewise, $LIG^{\perp}$ is LIG after excluding the effect by GDPG. Panel C of Table 8 shows the CSR results with the four factors, MKT, GDPG, SMB, and HML. The inclusion of GDPG makes the Fama and French factors (SMB and HML) insignificant. These results confirm Vassalou's (2003) results that when GDPG is present in the model, SMB and HML lose much of their ability to explain the cross-section of average stocks returns. However, when $GDPG^{\perp}$ is used instead of GDPG in the model, SMB and HML becomes significant (in Panel D). This means that much of the explanatory power of GDPG in Vassalou's results comes from LIG. When $LIG^{\perp}$ is used instead of GDPG in the four-factor model, $LIG^{\perp}$ is still significant, while SMB and HML are insignificant (in Panel E). Note that when LIG is present in the model, SMB and HML are insignificant (see Panels D and E of Table 4).

In fact, labor income is one component of GDP. The main part of the remaining component of GDP contains investment. Li et al. (2006) report that investment growth factor cannot capture the Fama and French factors. Thus, the above results are consistent with Li et al. (2006), since $GDPG^{\perp}$ contains mainly investment growth. Overall, our results show that labor income growth, rather than the whole GDP growth or the other components of GDP growth, is related to the Fama and French factors.

5. Robustness checks

This section provides a battery of robustness tests. These tests verify that our conclusions regarding future labor income growth risk are not driven by different estimation specifications; alternative base assets, labor income data, alternative horizons over which labor income growth is computed for constructing the economic tracking portfolio, and different frequency of data.

The above-mentioned estimation results for LIG could be sensitive to how to construct the tracking portfolio. Recall that for constructing the tracking portfolios, we use the following specifications: ten industry portfolios plus DEF and TERM as base assets, quarterly labor income data, discounted future labor income growth rates up to 12 quarters, and quarterly returns on the tracking portfolios.

It would be necessary, therefore, to perform a robustness test using various alternative specifications for constructing the tracking portfolios. Table 9 shows both CSR and GMM estimation results of our three-factor model in various alternative specifications.[23] We consider four different specifications: the Fama and French four portfolios (small growth, small value, large growth, and large value portfolios as base assets (in Panel A), monthly labor income data (in Panel B), discounted future labor income growth rates up to 4 years (or 16 quarters) (in Panel C), and monthly returns on the tracking portfolios (in Panel D). The reasons we use the Fama and French four portfolios as another set of base assets for the robustness check are that Breeden et al. (1989) argue that the same test assets should be used as a set of base assets in constructing tracking portfolios and that these four portfolios are also used as base assets in other studies (e.g., Malloy et al., 2008).[24]

The overall results are similar to those of the original specification for constructing the tracking portfolios. That is, in any specifications considered, the risk premium estimate on LIG is positive and statistically strongly significant. The *p*-value of Wald (SMB&HML) test is all greater than ten percent, which means that

[23] Since the consumption data are available in quarterly frequency, the two-factor model (MKT and LIG) without CONS is estimated (in Panel F) when returns on the tracking portfolio are given.

[24] Breeden et al.'s (1989) argument is as follows: Mathematically, $E(mR) = 1$ is equivalent to $E[proj(m|R)R] = 1$, where m denotes the stochastic discount factor, and R is the test assets.

Table 9

Cross-sectional regression and GMM estimation results of the three-factor model. The table reports the CSR estimation results (in the upper part) and the GMM estimation results (in the lower part) of our three-factor model. In Panel A, the Fama and French four portfolios used as alternative base assets are small growth, small value, large growth, and large value portfolios. In Panel B, *monthly* labor income (obtained from NIPA table 2.6) rather than quarterly labor income data is used in estimating LIG. In Panel C, future labor income growth from quarter t to $t + 16$ is used in estimating LIG. In Panel D, monthly LIG is used in the cross-sectional regression. The risk premiums associated with factors are estimated using Fama–MacBeth method. "JW corrected-t" is computed by using Jagannathan and Wang's (1998) correction for the errors-in-variables bias. The adjusted R^2 is computed by using Jagannathan and Wang (1996). The Wald (b) test is a joint significance test of the factor loadings in the pricing kernel. The HJ-distance is the Hansen–Jagannathan (1997) distance measure. The J-test is Hansen's (1982) test on the overidentifying restrictions of the model. The Wald (SMB&HML) statistic tests whether SMB and HML contain an incremental ability in pricing the test assets. The test statistics for Wald (b), HJ-distance, J-test, and Wald (SMB&HML) are computed through the GMM estimation. The sample period is from 1963:Q3 to 2007:Q4.

	Constant	CONS	MKT	LIG	Adj. R^2
Panel A: Fama–French four portfolios as alternative base assets					
Risk premium	−0.32	0.14	−0.99	0.32	0.80
t-value	−0.32	0.92	−0.90	3.83	
JW corrected-t	−0.27	0.81	−0.66	3.15	
	Tests				
	Wald (b)	HJ-distance	J-test	Wald (SMB&HML)	
Statistic	31.18	0.59	33.12	3.05	
p-value	0.00	0.05	0.04	0.22	
		CONS	MKT	LIG	
Panel B: Monthly labor income					
Risk premium	1.36	0.02	−0.40	0.81	0.72
t-value	1.53	0.13	−0.37	3.64	
JW corrected-t	0.94	0.08	−0.22	2.58	
	Tests				
	Wald (b)	HJ-distance	J-test	Wald (SMB&HML)	
Statistic	5.59	0.65	66.41	3.23	
p-value	0.13	0.00	0.00	0.20	
		CONS	MKT	LIG	
Panel C: Labor income growth over the 4 years (S = 16 quarters)					
Risk premium	2.25	−0.14	−2.34	0.84	0.78
t-value	2.53	−1.04	−1.96	3.84	
JW corrected-t	1.61	−0.68	−1.24	2.72	
	Tests				
	Wald (b)	HJ-distance	J-test	Wald (SMB&HML)	
Statistic	9.26	0.60	56.01	0.70	
p-value	0.03	0.03	0.00	0.70	
		MKT	LIG		
Panel D: Monthly frequency estimation					
Risk premium	−0.06	0.12	0.36		0.73
t-value	−0.12	0.24	3.83		
JW corrected-t	−0.08	0.16	2.90		
	Tests				
	Wald (b)	HJ-distance	J-test	Wald (SMB&HML)	
Statistic	8.24	0.32	46.74	4.08	
p-value	0.02	0.04	0.00	0.13	

when SMB and HML are added into the three-factor model, the coefficients on SMB and HML in the pricing kernel do not differ from zero. That is, their marginal explanatory power for the cross-sectional of average returns is insignificant. These results are consistent with Panel D of Table 4.

6. Conclusions

This paper proposes revisions in the expectation of future labor income growth as a macroeconomic state variable that is closely related to macroeconomic conditions and business cycle fluctuations and that may imply the nature of the economic risk captured by size and book-to-market equity. It then suggests a three-factor model that includes a factor related to this variable, along with the consumption growth factor and the market factor. This paper examines whether this future labor income growth factor captures the pricing abilities of the Fama–French factors in explaining the size and book-to-market effects. In order to obtain the risk factor that captures revisions in the expectation of future labor income growth, which are unobservable, we adopt the economic tracking portfolio approach introduced.

The CSR and GMM estimation results show that our three-factor model performs at least as well as the Fama and French model in explaining the cross-section of average returns of 25 size and book-to-market sorted test portfolios. In particular, our future labor income growth factor is consistently significantly priced in various model specifications. When the Fama and French factors are added into our three-factor model, they are no longer significant. This means that the future labor income growth factor is positively associated with the Fama–French factors and subsumes the explanatory power of these Fama–French factors in explaining the cross-section of stock returns. We interpret this positive association between the Fama–French factors and future labor income growth factor as suggesting that small firms and value stocks are more sensitive to shocks to the state of future labor income growth. Thus, our empirical results provide an economic explanation for the roles of the Fama–French factors in explaining equity returns: they are compensation for higher exposure to the risk related to revisions in the expectation of future labor income growth.

Since the results in this paper could be sensitive to the specification used for constructing the tracking portfolio, we perform robustness tests using various alternative specifications for constructing the tracking portfolio. However, the overall results are qualitatively the same. We also compare the performance of our three-factor model with other competing models in explaining the cross-section of average returns. We find that our proposed three-factor specification better captures cross-sectional variation in average returns than any of the competing asset pricing models considered.

Acknowledgements

This paper is based on third author's first chapter of doctoral dissertation at the Korea Advanced Institute of Science and Technology. The authors would like to thank Robert Dittmar, Mike Halling, Byoung Uk Kang, Jangkoo Kang, Kuan-Hui Lee, Jaewon Park, Oleg Rytchkov, Sun-Joong Yoon for many helpful and insightful comments. We also thank seminar participants at the European Finance Association Annual Meeting in Frankfurt; the Financial Management Association Annual Meeting in Reno; the 3rd Conference on Asia-Pacific Financial Markets in Seoul. We are particularly grateful to Ike Mathur (the editor) and an anonymous referee. Any remaining errors are our own responsibilities.

References

Abhyankar, A., Gonzalez, A., 2009. News and the cross-section of expected corporate bond returns. Journal of Banking and Finance 33, 996–1004.
Aretz, K., Bartram, S.M., Pope, P.F., 2010. Macroeconomic risks and characteristic-based factor models. Journal of Banking and Finance 34, 1383–1399.
Breeden, D.T., 1979. An intertemporal asset pricing model with stochastic consumption and investment opportunities. Journal of Financial Economics 7, 265–296.
Breeden, D.T., Gibbons, M.R., Litzenberger, R., 1989. Empirical tests of the consumption-oriented CAPM. Journal of Finance 44, 231–262.
Campbell, J.Y., 1996. Understanding risk and return. Journal of Political Economy 104, 298–345.
Campbell, J.Y., Cochrane, J.H., 2000. Explaining the poor performance of consumption-based asset pricing models. Journal of Finance 55, 2863–2878.
Campbell, J.Y., Shiller, R.J., 1988. The dividend-price ratio and expectations of future dividends and discount factors. Review of Financial Studies 1, 195–228.
Campbell, J.Y., Vuolteenaho, T., 2004. Bad beta, good beta. American Economic Review 94, 1249–1275.
Chan, K.C., Chen, N., 1991. Structural and return characteristics of small and large firms. Journal of Finance 46, 1467–1484.
Cochrane, J.H., 1996. A cross-sectional test of an investment-based asset pricing model. Journal of Political Economy 104, 572–621.
Epstein, L.G., Zin, S.E., 1989. Substitution, risk aversion, and the temporal behavior of consumption and asset returns: a theoretical framework. Econometrica 57, 937–968.
Epstein, L.G., Zin, S.E., 1991. Substitution, risk aversion, and the temporal behavior of consumption and asset returns: an empirical analysis. Journal of Political Economy 99, 263–286.
Fama, E.F., French, K.R., 1992. The cross-section of expected stock returns. Journal of Finance 47, 427–465.
Fama, E.F., French, K.R., 1993. Common risk factors in the returns on bonds and stocks. Journal of Financial Economics 33, 3–56.
Fama, E.F., French, K.R., 1995. Size and book-to-market factors in earnings and returns. Journal of Finance 50, 131–155.
Fama, E.F., French, K.R., 1996. Multifactor explanations of asset pricing anomalies. Journal of Finance 51, 55–184.
Fama, E.F., MacBeth, J., 1973. Risk, return, and equilibrium: empirical tests. Journal of Political Economy 71, 607–636.
Hahn, J., Lee, H., 2006. Yield spreads as alternative risk factors for size and book-to-market. Journal of Financial and Quantitative Analysis 41, 247–269.
Hansen, L.P., 1982. Large sample properties of generalized methods of moments estimators. Econometrica 50, 1029–1054.
Hansen, L.P., Jagannathan, R., 1997. Assessing specification errors in stochastic discount factor models. Journal of Finance 52, 557–590.
Jagannathan, R., Wang, Z., 1996. The conditional CAPM and the cross-section of expected returns. Journal of Finance 51, 815–849.
Jagannathan, R., Wang, Z., 1998. An asymptotic theory for estimating beta-pricing models using cross-sectional regression. Journal of Finance 53, 1285–1309.
Kim, D., 1995. The errors-in-variables problem in the cross-section of expected stock returns. Journal of Finance 50, 1605–1634.
Kim, D., 1997. A reexamination of size, book-to-market, and earnings-price in the cross-section of expected stock returns. Journal of Financial and Quantitative Analysis 32, 463–489.
Kim, D., 2010. Information uncertainty risk and seasonality in international stock markets. Asia-Pacific Journal of Financial Studies 39, 229–259.
Lamont, O., 2001. Economic tracking portfolios. Journal of Econometrics 105, 161–184.
Lettau, M., Ludvigson, S., 2001. Resurrecting the (C)CAPM: a cross-sectional test when risk premia are time-varying. Journal of Political Economy 109, 1238–1287.
Li, Q., Vassalou, M., Xing, Y., 2006. Sector investment growth rates and the cross section of equity returns. Journal of Business 79, 1637–1665.
Liew, J., Vassalou, M., 2000. Can book-to-market, size, and momentum be risk factors that predict economic growth? Journal of Financial Economics 57, 221–245.
Linter, J., 1965. Security prices, risk, and maximal gains from diversification. Journal of Finance 20, 587–615.
Lustig, H.N., Van Nieuwerburgh, S.G., 2008. The returns on human capital: good news on wall street is bad news on main street. Review of Financial Studies 21, 2097–2137.
Malloy, C.J., Moskowitz, T.J., Vissing-Jorgenson, A., 2009. Long-run stockholder consumption risk and asset returns. Journal of Finance 64, 2427–2479.
Moerman, G.A., van Dijk, M.A., 2010. Inflation risk and international asset returns. Journal of Banking and Finance 34, 840–855.
Newey, W., West, K., 1987a. Hypothesis testing with efficient method of moments. International Economic Review 28, 777–787.
Newey, W., West, K., 1987b. A simple positive-definite heteroskedasticity and autocorrelation consistent covariance matrix. Econometrica 55, 703–708.
Pantzalis, C., Park, J.C., 2009. Equity market valuation of human capital and stock returns. Journal of Banking and Finance 33, 1610–1623.
Petkova, R., 2006. Do the Fama–French factors proxy for innovations in predictive variables. Journal of Finance 61, 581–612.
Petkova, R., Zhang, L., 2005. Is value riskier than growth? Journal of Financial Economics 78, 187–202.
Shanken, J., 1992. On the estimation of beta-pricing models. Review of Financial Studies 5, 1–33.
Sharpe, W.F., 1964. Capital asset prices: a theory of market equilibrium under conditions of risk. Journal of Finance 19, 424–444.
Shiller, R.J., 1993. Aggregate income risk and hedging mechanism. Working paper, NBER.
Vassalou, M., 2003. News related to future GDP growth as a risk factor in equity returns. Journal of Financial Economics 68, 47–73.
Viale, A.M., Kolari, J.W., Fraser, D.R., 2009. Common risk factors in bank stocks. Journal of Banking and Finance 33, 464–472.
White, H., 1980. A heteroskedasticity-consistent covariance matrix estimator and direct test for heteroskedasticity. Econometrica 48, 817–838.