Labor Income, Relative Wealth Concerns, and the Cross-Section of Stock Returns*

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Abstract

The finance literature documents a relation between labor income and the cross-section of stock returns. One possible explanation for this is the hedging decisions of investors with relative wealth concerns. This implies a negative risk premium associated with stock returns correlated with local undiversifiable wealth, since investors are willing to pay more for stocks that help their hedging goals. We find evidence that is consistent with these regularities. In addition, we show that the effect varies across geographic areas depending on the size and variability of undiversifiable wealth, proxied by labor income.

I. Introduction

In this paper, we propose a channel that can explain the relationship between labor income and the cross-section of stock returns. In particular, the optimal hedging strategy of an investor with relative wealth concerns results in a multifactor equilibrium model in which the undiversified wealth of the investor's "peers" (for which, we argue, the component of labor income unrelated to stock market returns is a good proxy) is a negatively priced risk factor. We find strong empirical evidence in support of this channel.

Over the years, the finance literature has accumulated evidence of a connection between labor income and the cross-section of stock returns. Mayers (1972) is credited as the first to suggest the analysis of labor income as a measure of human capital in an asset pricing setting. In two influential papers, Campbell (1996) and Jagannathan and Wang (1996) use growth in labor income as a measure of the return on human capital. Their intuition is that human capital, a fundamental part of the economy's endowment, has been typically overlooked in the capital asset pricing model (CAPM). The inclusion of the return to human capital in empirical asset pricing models is able to explain a much higher portion of the cross-sectional variation in stock returns relative to the standard CAPM. Lettau and Ludvigson (2001a), (2001b) and Santos and Veronesi (2006) both introduce variables based on labor income into conditional asset pricing models and find that the explanatory power of the model increases substantially.

We consider a different channel. Our empirical evidence shows that labor income is related to the cross-section of stock returns through the hedging activity of investors with relative wealth concerns. This idea is based on the *KEEping up Pricing Model* (KEEPM) of relative wealth concerns developed in Gómez, Priestley, and Zapatero (2009). Investors hedge the risk that their reference group or "peers" will experience an income shock by investing in securities strongly correlated with the income of these peers. Equilibrium prices reflect the price pressure resulting from these hedging activities.

Relative wealth concerns implies restrictions on the relationship between human capital and stock returns not previously identified in the literature. First, the risk premium associated with the labor income factor is negative, since investors are willing to pay extra for securities that hedge this risk. Second, this relation must hold at the local level, since the main source of relative wealth concerns pertains to the surroundings of the investor.

We test the implications of the KEEPM using U.S. data. We begin using individual securities and study their relationship to the smallest unit for which we have disaggregated labor income, the state. We then undertake similar tests, using both individual securities and stock portfolios, at the US Census divisions level since this level of aggregation has been employed to examine local effects in the literature.[1] In addition, given the larger size

[1]The U.S. Census partitions the country in nine divisions (see Figure A1 in the Internet Appendix at www.jfqa.org).

of the divisions (some states have low Gross Domestic Product (GDP) and few stocks) we can perform further qualitative analysis. In particular, the model predicts that the hedging demand will be higher the higher the volatility of the factor investors want to hedge and the higher the relevance of that factor, as measured by the amount of undiversifiable labor income. The estimation of the model deep parameters shows evidence consistent with this prediction.

We compare the cross-sectional performance of our model with the performance of the CAPM and the 3-factor Fama and French (1992) model. In terms of pricing errors and R^2, our model performs much better than the CAPM and similarly to the three-factor model. Our risk factor is robust to the inclusion of the size and book-to-market factors from the Fama and French model. Finally, to double-check the local nature of the effect, we jointly test the local (state or division) factor and the aggregate (country) factor. We show that when we include both the country and the local (divisional or state) factors, both are priced and their risk premia are negative.

The literature has discussed two main sources of relative wealth concerns. On one hand, Keeping Up with the Joneses preferences, first introduced in the finance literature by Abel (1990) and further analyzed by Galí (1994); they show that in the absence of a market friction, optimal portfolio holdings are identical across investors and only market risk is priced. Brown, Ivković, Smith, and Weisbenner (2008) find that individual market participation increases with average community market participation. On the other hand, DeMarzo, Kaniel, and Kremer (2004) present a model of endogenous, price-driven relative wealth concerns; this idea is applied to technological investment and investment cycles in DeMarzo, Kaniel, and Kremer (2007) and to financial bubbles in DeMarzo, Kaniel, and Kremer (2008). Ravina (2007) presents evidence of this behavior using credit card data. Gómez (2007) analyzes its impact on portfolio choice. García and Strobl (2011) study the implications for information acquisition. Shemesh and Zapatero (2016) study its relationship with population density. Johnson (2012) finds that there exists a premium for stocks that hedge against income inequality.

Our paper is closely related to Korniotis (2008) who considers a consumption-based model of external habit formation as in Campbell and Cochrane (1999) for different partitions of the U.S. (the four U.S. census regions and eight Bureau of Economic Analysis (BEA) regions). These findings are in the spirit of Hong, Kubik, and Stein (2008). They show that the cross-section of stock returns depends on the census division where the headquarters of the firm are located. In this line of research, Korniotis and Kumar (2009) and Bernile, Korniotis, Kumar, and Wang (2015) show the connection between stock returns and local economic conditions.

Although we do not perform any direct test on portfolio holdings in this paper, the KEEPM yields partial equilibrium results that are consistent with those in the home bias literature that started with French and Poterba (1991).[2] Subsequently, a strand of the literature has shown a similar effect at the domestic level termed "home bias at home." Coval and Moskowitz (1999), for instance, study the investment behavior of money managers and observe that they favor (with respect to what would be optimal) local firms. Ivković and Weisbenner (2005) and Massa and Simonov (2006) show that U.S. and Swedish households, respectively, exhibit a strong preference for local investments. In our setting, investors in a given location (state or division) are willing to pay a premium for assets positively correlated with the divisional, non-diversifiable wealth. A related idea is the "familiarity" argument of Huberman (2001), who show that investors favor positions in local stocks. However, in our empirical work, we find that one factor that can explain the local bias is the correlation between labor income and security returns –regardless of the location of the firm.

The paper is organized as follows. We derive the KEEPM in Section II. Section III describes the data. In Section IV, we perform our baseline tests at the country level, pooling all securities in the tests. In Section V, we perform similar tests at the U.S. census divisional level. We then repeat the basic tests using aggregate labor income (instead of state labor income) as a proxy for undiversifiable wealth in Section VI. We close the paper with some conclusions. In addition, we have prepared an internet Appendix (we refer to it

[2]For a literature review of the home bias puzzle see Lewis (1999).

throughout the paper as "the Appendix").

II. The KEEPM

We consider the two main specifications discussed in the literature: exogenous and endogenous keeping up with the Joneses preferences. In both specifications, we assume a one-period economy with K geographical denominations. For the moment, let us assume that these denominations represent country divisions indexed by k (we will use both states and U.S. Census divisions in the empirical tests). In each division there is a local firm. At time $t = 0$, each firm issues one share that will yield a random payoff in time $t = 1$. We normalize the initial value of the firm to 1. Let r_k denote the random excess return on a share of firm k. The vector $r = (r_1, ..., r_k, ..., r_K)'$ has a joint distribution function $F(r)$, with mean return vector $E(r)$ and covariance matrix Ω. Firm shares can be freely traded across divisions. There is also a risk-free bond in zero net supply. Let R denote the return on the risk-free bond. Financial markets are complete. In each division there are two types of agents: "investors" and "workers," endowed with non-diversifiable stochastic local labor or entrepreneurial income.

We show in the Appendix that, whether endogenous or exogenous, relative wealth concerns and non-diversifiable income implies the following optimal portfolio for the representative investor in division k:

$$x_k^* \quad = \quad \theta_k b_k X_k^w + \tau_k \Omega^{-1} E(r), \tag{1}$$

where X_k^w represents a mimicking portfolio that maps the workers endowment return onto the investment opportunity set; θ_k denotes the the relative wealth at $t = 0$ of the division's workers as a proportion of the total division's wealth. The parameters b and τ represent the portfolio bias and the risk-tolerance coefficient, respectively, with values:

JONESES	b	τ
Exogenous	$\frac{\gamma}{1-\gamma}$	$\frac{1}{\alpha(1-\gamma)}$
Endogenous	$\frac{\alpha-1}{\alpha}$	$\frac{1}{\alpha}$

Notice that, given these definitions, there will exist a bias in portfolio holdings towards the Joneses portfolio (hence, consumption) only if $0 < \gamma < 1$, in the exogenous specification, and $\alpha > 1$, in the endogenous specification.[3]

Market clearing in financial markets at time $t = 0$ requires that $\sum_k \omega_k x_k^* = x_M$, with x_M the market portfolio, with excess return r_M, and $\omega_k = c_k^0 / \sum_k c_k^0$. Spot market clearing at time $t = 1$ implies that workers consume the proceedings of their (non-tradable) endowment, w, and investors the return on their portfolios, c. We regress the workers non-diversifiable wealth return, $r_k^w = r' X^w$, onto the country market portfolio excess return:

[3]The constraint on $\alpha > 1$ is already present in DeMarzo, Kaniel, and Kremer (2004).

$$(2) \qquad r_k^w = \beta_k \, r_M + r_k^F.$$

Portfolio $\beta_k \, x_M$ represents the projection of the workers income onto the security market line spanned by the aggregate market portfolio x_M. Define the portfolio $F_k \equiv X_k^w - \beta_k \, x_M$ as an orthogonal factor portfolio with return $r_k^F = r' F_k$ and mean return μ_k^F. After these definitions, the workers' portfolio can be expressed as a linear combination of the market portfolio and a zero-beta (orthogonal) portfolio: $X_k^w = F_k + \beta_k x_M$. We replace X_k^w in equation (1):

$$x_k^* \;=\; \theta_k \, b_k \, F_k + \theta_k \, b_k \, \beta_k \, x_M + \tau_k \, \Omega^{-1} E(r).$$

This portfolio has three components. Portfolio F_k is division-specific and can be interpreted as a *hedge portfolio*: portfolio F_k hedges investors from the risk involved in keeping up with the local non-diversifiable Joneses risk. Given the orthogonality conditions, this portfolio plays the role of a division-specific, zero-beta asset. The projection component, $\beta_k x_M$, corresponds to that part of the workers wage income perfectly correlated with the country market portfolio. The standard component, $\Omega^{-1}E(r)$, is the highest global Sharpe ratio portfolio and it is common across divisions.

We define the coefficient H as the inverse of the risk-tolerance coefficient $H^{-1} = \sum_k \omega_k \tau_k$. After imposing market clearing, we solve for equilibrium expected returns:

$$E(r) \;=\; H\,\Omega \left[\left(1 - \sum_{k=1}^{K} \omega_k \theta_k \, b_k \, \beta_k \right) x_M - \sum_{k=1}^{K} \omega_k \theta_k \, b_k \, F_k \right]. \tag{3}$$

Define the matrix $\boldsymbol{F}$ of dimension $N \times (K+1)$ as the column juxtaposition of the market portfolio and the orthogonal portfolios, $\boldsymbol{F} \equiv (x_M, F_1, ..., F_k, ..., F_K)$. Let $\boldsymbol{r^F} \equiv (r_M, r_1^F, ..., r_k^F, ..., r_K^F)$ denote the vector of factor returns. Additionally, define the wealth vector as

$$\boldsymbol{W} \;\equiv\; H \left(1 - \sum_{k=1}^{K} \omega_k \theta_k \, b_k \, \beta_1, -\omega_1 \theta_1 \, b_1, ..., -\omega_k \theta_k \, b_k, ..., -\omega_K \theta_K \, b_K \right)'.$$

Given these definitions, the equilibrium condition of equation (3) can be re-written as $E(r) = \Omega\,\boldsymbol{FW}$. Pre-multiplying both terms of the previous equation by the transpose of matrix $\boldsymbol{F}$ we obtain the equilibrium condition for the vector of prices of risk, $\boldsymbol{\lambda} \equiv (\lambda^M, \lambda^1, ..., \lambda^k, ..., \lambda^K)$, with the market risk premium, λ^M, as the first component. Thus, $\boldsymbol{\lambda} = \boldsymbol{F}'\Omega\,\boldsymbol{F}\,\boldsymbol{W}$, where $\boldsymbol{F}'\Omega\,\boldsymbol{F}$ is a matrix of dimension $(K+1) \times (K+1)$ whose first column (row) includes the market return volatility and a vector of K zeros and the remaining elements are the covariances between F_k and $F_{k'}$ The expected risk premia on the market and the zero-beta portfolios will be:

$$\lambda^M = H\left(1 - \sum_{k=1}^{K} \omega_k \theta_k \, b_k \, \beta_k\right) \sigma_M^2, \tag{4}$$

$$\lambda^k = -H\left(\omega_k \theta_k \, b_k \operatorname{var}(r_k^F) + \sum_{k' \neq k} \omega_{k'} \theta_{k'} \, b_{k'} \operatorname{cov}(r_k^F, r_{k'}^F)\right). \tag{5}$$

The market portfolio, x_M, is partially correlated with each division's non-diversifiable risk. This correlation is captured by the coefficient β_k and offers partial hedging against deviations from the local Joneses (in case $\theta b > 0$). Therefore, the equilibrium price of risk for the country market risk factor, λ^M, is different from the symmetric equilibrium. The parenthesis in equation (4), which in the case of a symmetric equilibrium would be 1, captures the net price of risk on the aggregate market risk factor, after discounting the (capitalization weighted) Joneses hedging effect. If the weighted value of the betas is higher than the country market beta (i.e., 1), the market price of risk could turn negative: if the hedging properties of the market portfolio against Joneses deviations outweigh the compensation for systematic risk, the *net* expected market price of risk becomes negative.

More importantly, if there is a relative wealth concern ($b > 0$) in the economy and workers income is not diversifiable ($\theta > 0$), there should be K additional risk factors (one per division) to the market risk factor. Regarding their sign, the model predicts that if $\text{cov}(r_k^F, r_{k'}^F) > 0$ for all k, k', then every λ^k will be negative.[4] To understand this result, suppose for the moment that the zero-beta portfolios were orthogonal ($\text{cov}(r_k^F, r_{k'}^F) = 0$) for all k, k'. Then, the price of risk would be strictly negative: An asset that has positive covariance with portfolio F_k will hedge the investor in division k from the risk of deviating from the non-diversifiable (local) income of the Joneses. This investor will be willing to pay a higher price for the asset thus yielding a lower return. In equilibrium, the price of risk for F_k would be, in absolute terms, increasing in b_k and the volatility of the hedge portfolio. If the covariance across zero-beta portfolios is positive, this just increases the absolute value of the negative prices of risk for every division's hedge portfolio. Solving for $\boldsymbol{W}$ we obtain:

$$E(r) = \boldsymbol{\beta}^{\boldsymbol{F}} \boldsymbol{\lambda}, \tag{6}$$

where $\boldsymbol{\beta}^{\boldsymbol{F}} = \Omega \boldsymbol{F} (\boldsymbol{F}'\Omega \boldsymbol{F})^{-1}$ denotes the $K \times (K+1)$, in general, for N assets,

[4]Notice that this is a sufficient condition satisfied by our data in Panel B of Table 4.

$N \times (K+1)$, matrix of betas, with the first column as the market betas for all assets.

We call this pricing model that captures the equilibrium implications of relative wealth concerns, both under the exogenous and endogenous specifications KEEPM, ("KEEping up Pricing Model"). In the following sections, we test the models' restrictions in equations (4), (5), and (6).

III. Data Description

To construct the risk factors that will proxy relative wealth concerns, we need to make some assumption regarding the geographical dimension of "the peers."For example, should they be defined at the city, state, division, region or national level? Arguably, the relevance of keeping up with the Joneses should be higher (larger γ in the model) at the state level, the smallest unit for which we have data on labor income, than at the divisional or national levels. Therefore, from Compustat, we obtain annual information on headquarter location for the period 1963 to 2011. Consistently with previous studies, we exclude Hawaii and Alaska to avoid biases in our results. Using this information, we obtain stock returns for all NYSE, AMEX and NASDAQ stocks from CRSP for 1960Q1 to 2011Q4.

For each stock, we proxy local non-diversifiable wealth using personal income data from the BEA corresponding to the state where the company's headquarters are located.

Following Santos and Veronesi (2006), we calculate the return on personal income per capita in quarter t by dividing the difference in personal income between quarter t and quarter $t-1$ by the personal income in quarter $t-1$, all per capita.

Following equation (2) in the model, for each state s we regress the return on state level personal income per capita on the CRSP aggregate stock market excess return and use the residuals from this regression as the orthogonal return on state labor income, denoted by r_s^F. As a robustness test, we replace the state labor income with the divisional labor income. Following the same orthogonalization procedure, we obtain for each division k the time series of orthogonal divisional labor income return, denoted by r_k^F. Finally, in order to compare the local versus country effect of the Joneses behavior, we calculate the U.S. country labor income per capita. The corresponding orthogonal country labor income return is denoted by r^C.

Regarding the test assets, we use individual assets and assets sorted into portfolios. Using individual assets, we test in the first place the cross-sectional predictions of the model at the country level. This requires using all U.S. individual stocks jointly, regardless of their headquarters location, as test assets. This approach presumes that the price of risk associated to the non-diversifiable labor income risk is the same across all states and divisions.

To compare the performance of our model with other standard models in the literature (notably, the CAPM and 3-factor Fama and French (1992) model), we replace the individual stocks with portfolios. At the same time, we construct factor mimicking portfolios for the orthogonal labor income risk, both at the aggregate level and in the divisional tests. In addition to the local risk factors, we also require the excess return on the aggregate stock market portfolio (ERM), as proxied by the CRSP aggregate index, the small minus big market capitalization portfolio (SMB) and the high minus low book to market portfolio (HML). All these portfolios are taken from the web site of Kenneth French. The quarterly premia on ERM, SMB and HML are 1.33%, 0.85%, and 1.32%, respectively, over the sample period.

IV. Country Level Tests

Our first test of the model assumes that the price of risk for the local orthogonal labor income risk is unique across states and divisions. This is a strong assumption that we will relax in the following section where the tests will be conducted division by division. The country level tests in this section offer the first evidence in favor of the model's main prediction: namely, that there exists a negative price of risk on the orthogonal state labor income return. We also compare the cross-sectional performance of our model relative to the performance of other established asset pricing models in the literature.

Starting in 1960, we use five years of quarterly data and regress the return on every individual stock i in the U.S. on a constant, the orthogonal state labor income return, $r^F_{s,t}$, and the CRSP aggregate stock market excess return, $r_{\mathrm{RM},t}$:[5]

$$r_{i,t} \;=\; \alpha_i + \beta_i^F r^F_{s,t} + \beta_i^{\mathrm{RM}} r_{\mathrm{RM},t} + u_{i,t}. \tag{7}$$

We then add one quarter of data and re-estimate. We keep adding one quarter of returns and re-estimating the orthogonal beta and the market beta until we have 36 quarters. After that point, every time a new quarter of data is added, the first quarter is removed and the process is repeated. The time series of quarterly estimated rolling betas starts in 1965Q1 and ends in 2011Q4. We use this time series to run cross-sectional regressions, quarter by quarter, to estimate the price of risk on the state orthogonal labor income factor:[6]

$$r_i \;=\; \lambda_t^0 + \lambda_t^F \hat{\beta}_i^F + \lambda_t^{\mathrm{RM}} \hat{\beta}_i^{\mathrm{RM}} + \xi_i. \tag{8}$$

Table 1 presents the time series averages of the intercept λ^0, and the prices of risk, λ^F and λ^{RM} where absolute t-values are reported in parenthesis. As predicted by the model, the average price of risk on the orthogonal labor income risk is negative and strongly significant

[5] We assume that a firm that is headquartered in state s in 1963 is headquartered in that state in 1960, 1961 and 1962.

[6] All cross-sectional results are qualitatively analogous when the prices of risk are estimated with respect to the one-year lagged betas.

with an absolute t-value equal to 2.27.[7] The size of the orthogonal risk premium is economically significant at -0.198. This implies that a stock with a unit beta on the orthogonal local labor income factor has a quarterly return twenty basis points lower than a stock with a zero beta. This lower return reflects the fact that a stock with a unit beta is a good hedge for orthogonal local labor income and its price has been pushed up, and hence returns are lower. The market price of risk, λ^{RM}, is 1.4% per quarter implying an annual equity market risk premium of around 5.5%. However, the intercept at 1.9% per quarter, which should be equal to the risk free rate of return, is large suggesting some model misspecification. One possible explanation for this is the restriction that the price of risk associated with the local relative wealth concerns is forced to be the same for every stock. Another explanation is that there are missing risk factors.[8]

The findings in Table 1 illustrate the importance of labor income in explaining the cross-section of individual stocks returns. The negative estimate on the price of risk associated with this local factor suggests that stocks that have a hedging potential for

[7] Recall that the estimate sign on λ_t^F should be negative. Therefore the test is one-sided.

[8] The positive and statistically significant intercept may be capturing risk resulting from other factors that might have a positive price. For example, hedging demands from peer-dependent preferences related to the agents concern for status (as in Roussanov (2010)). To assess whether the negative prices of risk on the local risk factor are affected by this, we have re-estimated the cross-sectional regressions omitting the intercept. We find that this has no material impact on the size, sign or statistical significance of the prices of risk on the local factors in the KEEPM model. Results are available from the authors. We thank the referee for suggesting this test.

investors have lower expected returns, since investors are willing to pay a premium to hold these stocks.

It is standard in empirical asset pricing tests to use portfolios of stocks as test assets in order to reduce the errors in variables problem that plagues the two-step Fama and MacBeth (1973) methodology. In addition, it is common to use factor mimicking portfolios to proxy risk factors in order to be able to interpret the estimated prices of risk in terms of returns (risk premia). Furthermore, model performance that focuses on pricing errors is easier to undertake with the use of well diversified portfolios. On the other hand this approach also has well-known problems, as documented in Daniel and Titman (2011) and Lewellen, Nagel, and Shanken (2010), for example. We perform these tests for robustness purposes and discuss them in the Appendix. The results corroborate the findings of this section.

V. Tests Per Division

We now focus on the divisional level. The objective is to test if the intensity of keeping up with the Joneses varies across U.S. divisions and whether this is reflected in the size of the orthogonal labor income price of risk in a way consistent with the predictions of the model. Every state belongs to one of the nine Census Bureau divisions: West South Central (WS), Pacific (PA), East South Central (ES), Mountain (MO), East North Central

(EN), South Atlantic (SA), West North Central (WN), Middle Atlantic (MA), and New England (NE).

A. Individual Stocks

Stocks are first sorted into divisions according to the location of the company's headquarters. We then follow the procedure explained in Section IV and estimate, for each stock in the division, the betas with respect to the orthogonal state labor income and the U.S. stock market beta from equation (7). We then run Fama-MacBeth cross-sectional regressions at each quarter t from 1965Q1 through 2011Q4 using as dependent variable the stock return, and as independent variables the estimated orthogonal, $\hat{\beta}_i^F$, and market, $\hat{\beta}_i^{\mathrm{ERM}}$, betas. The only difference with respect to the cross-sectional tests in the previous section is that we use only stocks headquartered within each division. In particular, for each division k we run:

$$r_{i,k} = \lambda_{t,k}^0 + \lambda_{t,k}^F \hat{\beta}_i^F + \lambda_{t,k}^{\mathrm{ERM}} \hat{\beta}_i^{\mathrm{ERM}} + \xi_{i,k}. \tag{9}$$

Panel A of Table 2 reports the intercept, λ_k^0, and the average prices of risk, λ_k^F and λ_k^{ERM} for each division k. Considering a one sided test, all the orthogonal prices of risk are negative and significant at the 5% level with the exception of MA, only marginally significant at the 10%, and WN. In the cases of SA, PA, ES and MO, the price of risk is statistical significant at the 2.5% level, with SA and PA significant at the 0.5% level. In terms of size, there is a wide discrepancy across divisions: from the smallest in absolute value in the case of MA (-0.165) to the largest corresponding to MO (-0.324).

As a robustness test, we replace the orthogonal state labor income r_s^F with the orthogonal divisional labor income, r_k^F. We then run, for each division, the time series regression of equation (7) for every stock in the U.S., regardless of the location of the company's headquarters. We estimate the corresponding betas with respect to the orthogonal divisional labor income and the market. These betas replace $\hat{\beta}_i^F$ and $\hat{\beta}_i^{\mathrm{ERM}}$, respectively, in equation (9). The average prices of risk λ_k^0, λ_k^F and λ_k^{ERM} for each division k are reported in Panel B of Table 2. The prices of risk on the orthogonal factors are, overall, very similar to those of Panel A. In the case of MA, ES and MO the price of risk increases marginally in absolute terms (more notably in ES), whereas in EN and WN it decreases, remaining practically the same in the other divisions.

The differences in the prices of risk of Panels A and B can be understood as follows. Arguably, as reasoned in Section III, the relevance of keeping up with the Joneses should be higher (larger γ in the model) at the state level. On the other hand, as it is clear from equation (5), the size of the orthogonal price of risk depends on the volatility of the "local" (i.e. divisional) orthogonal factor and the weighted covariance with the orthogonal factors from other divisions. Insofar as these factors are correlated, holding or shorting stocks from other divisions may affect the average orthogonal price of risk. These two effects partially compensate each other. A comparison of Panels A and B in Table 2 reveals that the net effect varies across divisions although it is, on average, very small. These results suggest that most of the hedging against the risk of deviating from the local Joneses consumption comes from the stocks of firms that are located closer to the source of non-diversifiable labor income, consistent with the documented home-bias at home phenomenon in U.S. portfolio holdings (Coval and Moskowitz (1999) and Brown at al. (2008)).

B. Portfolios

We construct a factor mimicking portfolio for the orthogonal state labor income risk in each division: Each year t, we sort stocks within each division into three equally-weighted portfolios, from the first quarter of 1965 to the final quarter of 2011, based on the coefficient on orthogonal labor income, $\hat{\beta}_i^F$, estimated until year $t-1$. The returns of the factor mimicking portfolio are computed as the returns of the portfolio (P1) formed by the stocks with the highest one third of coefficient estimates minus the returns on the portfolio (P3) formed by the stocks with the lowest one third of coefficient estimates. We represent by $r_{k,t}^{\mathrm{FM}}$ the time-series return on the state factor mimicking portfolio in division k.

Similarly, to generate the beta-sorted test portfolios we repeat the procedure discussed above and construct ten equally weighted portfolios per division.[9] We calculate excess returns on all the test portfolios by subtracting the 1-month T-bill rate from the actual returns.

Panel A of Table 3 reports the average return spread between portfolio P1 and portfolio P3 (alternatively, portfolio P10) for each division. All spreads are negative. As in Panel A of Table 2, the spreads P1–P3 are not uniform in size across divisions. They range from −0.429 in NE to −2.081 in PA. In four out of the nine divisions (MA, PA, ES, and WS) the spread is different from 0 at least at the 5% confidence level (at the 0.5% level in the case of PA and WS). When we analyze P1–P10, the spread increases in (absolute) size for all divisions except PA, where it marginally decreases, and ES.

In Panel B of Table 3, we recalculate the spreads using the divisional labor income return in each division. We use all stocks regardless of their headquarter's location. The effect varies from division to division. Looking first at P1–P3, compared to Panel A, the spreads increase in all divisions except in PA, ES, and WS, where they decrease, although they still remain strongly significant. All spreads are now statistically significant at least at the 5% level, which is probably due to the fact that the portfolios contain more stocks. When we compare P1–P10 with P1–P3 in Panel B of Table 3, the spreads increase in

[9] All the results presented in the paper are generally robust to the use of market capitalization weighted portfolios.

(absolute) size in all divisions.

We report the excess returns for each portfolio in Panel C of Table 3. The portfolios in this panel are created by sorting stocks within each division with respect to the coefficient $\hat{\beta}_i^F$ estimated with respect to the orthogonal state labor income return. Virtually all returns are strongly different from zero and they tend to increase in size as we move from P1 to P10 indicating that there is a reasonable spread in returns driven by the loadings on the mimicking portfolio.

Panel D of Table 3 reports, for each division k and each portfolio p, the coefficient with respect to the orthogonal state labor income factor mimicking portfolio, $r_{k,t}^{\mathrm{FM}}$, from:

$$r_{p,k,t} = \alpha_{p,k} + \beta_{p,k}^{\mathrm{FM}} r_{k,t}^{\mathrm{FM}} + \beta_{p,k}^{\mathrm{ERM}} r_{\mathrm{ERM},t} + u_{p,k,t}. \tag{10}$$

Most of the estimated coefficients are statistically significant. They decrease is size as we move from portfolios with higher covariance with the orthogonal state labor income (P1) to portfolios with lower covariance (P10). This, together with the negative price of risk on the orthogonal risk factor reported in Table 2, implies that portfolios more correlated with the orthogonal state labor income carry a lower expected return.

We study now the cross-sectional performance division by division. In each division k, we run the following contemporaneous regression each quarter t from 1965Q1 until

2011Q4 for the ten portfolios sorted by the orthogonal state labor income beta of Table 3:

$$r_{p,k} = \lambda^0_{t,k} + \lambda^{\text{FM}}_{t,k}\hat{\beta}^{\text{FM}}_{p,k} + \lambda^{\text{ERM}}_{t,k}\hat{\beta}^{\text{ERM}}_{p,k} + \xi_{p,k}, \tag{11}$$

The results are reported in Panel A of Table 4. The estimated intercept, λ^0_k, is not statistically different from zero in any division. Qualitatively, the estimated prices of risk λ^{FM}_k are very similar to the P1–P3 spreads reported in Table 3 Panel A. They are all negative, and range from -0.101 in EN to -1.812 in PA. Five out of the nine risk premia are significant at least at the 5% confidence level (at the 0.5% level in the case of PA and WS).

The last three columns in Panel A of Table 4 report, for each division, the cross-sectional regression adjusted $\overline{R}^2$, the average pricing errors and the test of whether the pricing errors are jointly zero. $\overline{R}^2$ ranges from 17% for EN to 92% for WS.[10] In all divisions with high factor mimicking variance (PA, ES, WS, and MO) the cross-sectional power of the test is above 60%. The pricing errors, defined as the difference between the actual portfolio return and the expected return, are small relative to the average portfolio return reported in Panel C of Table 3. In all cases the test rejects the null hypothesis that

[10]Following Jagannathan and Wang (1996) and Lettau and Ludvigson (2001b), we calculate R^2 as $\left[var_c(\overline{r}_p) - var_c(\overline{\xi}_p)\right]/var_c(\overline{r}_p)$, where var_c is the cross-sectional variance, $\overline{r}_p$ is the average return and $\overline{\xi}_p$ is the average residual. $\overline{R}^2$ is the adjusted R^2.

these pricing errors are different from zero.[11] It is worth noting that when we performed the same test at the aggregate level for all U.S. beta-sorted portfolios simultaneously (reported in the Appendix) the null hypothesis could not be rejected. We interpret this evidence as support for the KEEPM model at the local level where we allow for the price of risk on the orthogonal factor mimicking portfolio to vary across divisions.[12]

The variation in the size of prices of risk is consistent with the predictions of the model. In particular, observe that if we ignore the covariance terms, the value of the price of risk on the orthogonal labor income return is, according to equation (5), a function of three factors. First, the proportion of local non-diversifiable wealth in the division, $\omega_k\theta_k$; second, the Joneses preference parameter, γ_k; third, the variance of the orthogonal labor income return, $\text{var}(r_k^F)$.

We test empirically the model's prediction on the Joneses parameter, γ, in Subsection C. According to the BEA (a map is included in the Appendix, Figure A2) there is a high concentration of non-diversifiable wealth (proxied in our tests by personal income)

[11]This is a Chi-sq test given as $\widehat{\alpha}' cov(\widehat{\alpha})^{-1}\widehat{\alpha}$, where $\widehat{\alpha}$ is the vector of average pricing errors across the forty-five portfolios and cov is the covariance matrix of the pricing errors.

[12]To check the robustness of our results to the homecedasticity assumption implicit in the OLS cross sectional estimates, Table 4 in the Appendix reports the General Method of Moments (GMM) estimates of the prices of risk and their factor loadings in an approximate linear stochastic discount factor derived from the KEEPM equilibrium conditions. The results are very similar to those reported in Table 4 Panel A; in some divisions, like MA and EN, even stronger. We thank the editor and the referee for suggesting this robustness test.

in certain states and divisions. PA, MA, EN, SA, and WS are the divisions with higher concentration and MO, WN, NE, and ES are the divisions with lower concentration.

Regarding the effect of volatility of labor risk factors, Panel B of Table 4 shows, on the diagonal, the variance of the divisional factor mimicking portfolios. There is wide heterogeneity. Divisions with high factor volatility like PA (0.86%), ES (0.83%), WS (0.83%) and MO (0.89%) exhibit the largest (absolute) orthogonal prices of risk in Panel A of Table 4. Within these divisions, PA (−1.812%) and WS (−1.805%) have the absolute largest premia, and they are both strongly significant at the 0.5% confidence level –both divisions comprise states with a high concentration of personal income.

In contrast, ES (−1.456%) and MO (−1.193%) have relatively smaller premia, significant at the 5% only in the case of MO. Both divisions include states with a low concentration of personal income. Among the rest of divisions, some of them have either low factor volatility like NE (0.33%) and EN (0.22%) or a small concentration of personal income like WN. MA (0.37%) and SA (0.59%) have relatively low factor volatility but both divisions include states with high concentration of personal income. This may explain why their (absolute) prices of risk are relatively large in comparison with other divisions with similar factor volatility but lower concentration of personal income.

Finally, in Panel C of Table 4, we repeat the cross-sectional tests in equation (11) including the size, λ_k^{SMB}, and book-to-market, λ_k^{HML}, risk factors in each division:

$$r_{p,k} \;=\; \lambda^{0}_{t,k} + \lambda^{\text{FM}}_{t,k}\hat{\beta}^{\text{FM}}_{p,k} + \lambda^{\text{ERM}}_{t,k}\hat{\beta}^{\text{ERM}}_{p,k} + \lambda^{\text{SMB}}_{t,k}\hat{\beta}^{\text{SMB}}_{p,k} + \lambda^{\text{HML}}_{t,k}\hat{\beta}^{\text{HML}}_{p,k} + \xi_{p,k}.$$

The prices of risk of the orthogonal factor mimicking portfolios λ^{FM}_{k} are very similar to the estimates from the KEEPM reported in Panel A of Table 4, both in size and significance, perhaps with the exception of ES that increases (in absolute value) from -1.456 to -2.771 and turns marginally significant at the 10% level. The estimate prices of risk the 2 Fama and French factors are generally not statistically significant.

The average pricing errors are similar in size to those reported in Panel A of Table 4 for the KEEPM while the test of whether the joint pricing errors are statistically different from zero is rejected in all divisions except MA. We therefore conclude that the orthogonal state labor income factor that captures the risk of deviating form the Joneses consumption in each division is robust to the inclusion of the Fama and French risk factors. Moreover, after analyzing the pricing errors and the explanatory power of the tests, the cross-sectional performance of the KEEPM, division by division, improves relative to the aggregate (all U.S. stocks simultaneously) performance and does not improve in a significant way when we introduce the Fama and French risk factors. Overall, the results reported in Table 4 provide support for the KEEPM but also show that Keeping up with the Joneses behavior is not uniform across divisions, which is consistent with the model.

C. Estimation of the Joneses Parameter

The equilibrium conditions in equations (4) and (5) link explicitly the prices of risk to the deep parameters of the model. We derive the following orthogonality conditions from these equations:

$$0 = \lambda^{\text{ERM}} - H\left(1 - b\sum_{k=1}^{K}\omega_k\theta_k\,\beta_k\right)\left(r_M - \text{E}(r_M)\right)^2, \tag{12}$$

$$\begin{aligned} 0 &= \lambda_k^{\text{FM}} + H\,b\left(\omega_k\theta_k\,\left(r_k^{\text{FM}} - \text{E}(r_k^{\text{FM}})\right)^2\right. \\ &\quad \left. + \sum_{k'\neq k}\omega_{k'}\theta_{k'}\,\left(r_k^{\text{FM}} - \text{E}(r_k^{\text{FM}})\right)\left(r_{k'}^{\text{FM}} - \text{E}(r_{k'}^{\text{FM}})\right)\right), \end{aligned} \tag{13}$$

for each division $k = 1, 2, ..., K$. Ideally, we would like to estimate every divisional Joneses parameter, γ_k. The system, however, is not uniquely determined when we allow this parameter to vary across divisions. Hence, we assume a common γ across divisions. This implies $b = \gamma/(1-\gamma)$.[13] r_M and r_k^{FM} denote, respectively, the time series of the U.S. market return and the return on the factor mimicking portfolio from division k, from the first quarter of 1965 to the final quarter of 2011. We take the estimates of the price of risk for

[13] We only present the results for the exogenous specification. The estimation does not converge when we try to estimate the endogenous version of the model.

each division, λ_k^{FM}, from Table 4 Panel A. We proxy for θ_k using the time series of divisional personal income as a proportion of the divisional GDP. To proxy ω_k we divide each quarter the market capitalization of all stocks in the division by the aggregate market capitalization. The GMM methodology outlined in Hansen (1982) provides a natural way to estimate the deep parameters of our model. The system of equations (12) and (13)at the divisional level ($K = 9$) involves $N = 10$ moment conditions. We assume different values of the aggregate risk aversion coefficient H and estimate $L = 2$ parameters: the parameter γ and the market price of risk, λ^{ERM}.[14]

Table 5 presents the results for a two-step GMM estimation. The initial value for the Joneses parameter is $\gamma = 0.1$. The results are robust to alternative initial values. The estimate of γ is, both economically (the model predicts it should be bigger than zero and smaller than one) and statistically significant at the 1% level in all cases. There is a clear inverse relation between γ and H, supported by the model in the definition of λ^k in equation (5). Intuitively, the absolute size of the price of risk for the non-diversifiable income risk in a given division depends, directly, on the aggregate risk aversion coefficient, H, and the Joneses parameter, γ. Given the estimated prices of risk from Table 4 Panel A, a higher assumed value for H results, consequently, in a lower estimated value for γ.

[14]The ability of the model to price the assets is assessed by testing that the orthogonality conditions, which follow a $\chi^2(N - L)$ distribution, are zero. This is known as Hansen's J-test.

Notice that the estimate of λ^{ERM} is negative. This is consistent with the equilibrium equation for the market price of risk, λ^M, in the equilibrium condition of equation (4). In particular, if $b \sum_{k=1}^{K} \omega_k \theta_k \, \beta_k > 1$, the model predicts that the hedging property of the market portfolio $vis - \grave{a} - vis$ the risk of deviating from the Joneses portfolio outweighs the traditional positive market risk-reward mechanism. The "net" result is a negative market premium. Since γ decreases with H, the average value of $b \sum_{k=1}^{K} \omega_k \theta_k \, \beta_k$ ranges from 8.06 for $H = 1$ to 1.44 for $H = 6$, higher than 1 in all cases. This also explains why λ^M decreases in absolute value with H in Table 5. For higher H, the model implies a lower estimate of γ. Thus, hedging Joneses risk becomes less relevant. The market risk premium, although negative and statistically significant at the 1% level for all the values of H, becomes smaller in absolute terms.

Since γ is assumed to be constant, all variation in the estimated prices of risk reported in Panel A of Table 4 must come, according to equations (4) and (5), from the interaction of the percentage of divisional non-diversifiable wealth (relative to total country wealth), $\omega_k \theta_k$, and the volatility of the orthogonal risk factors. In other words, the GMM test offers an explanation for the variation in the prices of risk across divisions based on exogenous Joneses risk-hedging consistent with the predictions of the KEEPM model.

Whilst we obtain sensible estimates of the model's parameters, the J-test rejects the model, like in Korniotis (2008). We cannot rule out the possibility that this rejection is the result of forcing the parameter γ to be the same across divisions. As we have shown in the previous subsection, there is evidence that the factor is local. Therefore, the imposition that the relative wealth concerns parameter is the same across regions would seem to be too restrictive and leads to a rejection of the model.

VI. Country-Wide Orthogonal Income

We now study the cross-sectional performance of the KEEPM when the orthogonal state labor income return is replaced with the orthogonal U.S. country labor income return. Gómez, Priestley, and Zapatero (2009) show that the orthogonal U.S. country labor income return carries a negative price of risk. The evidence reported so far in this paper points in the direction of a local hedging demand that varies across divisions. Our objective is to compare the divisional and country performance of the KEEPM and test whether the variation in the prices of risk across divisions persist after considering jointly local and country risk factors. The results of these tests are presented in Table 6.

We denote by r_t^C the orthogonal country labor income return. We follow the procedure of in Section IV. For each individual stock i in the U.S. we estimate the rolling betas with respect to the orthogonal country labor income return and the U.S. stock market return:

$$r_{i,t} \quad = \quad \alpha_i + \beta_i^C r_t^C + \beta_i^{\mathrm{ERM}} r_{\mathrm{ERM},t} + u_{i,t}.$$

The slope coefficients $\hat{\beta}_i^C$ and $\hat{\beta}_i^{\mathrm{ERM}}$ are estimated for every stock in the U.S. The time series of quarterly estimated betas starts in 1965Q1 and ends in 2011Q4. We then run the Fama-MacBeth cross-sectional regressions at each period t:

$$r_i \quad = \quad \lambda_t^0 + \lambda_t^C \hat{\beta}_i^C + \lambda_t^{\mathrm{ERM}} \hat{\beta}_i^{\mathrm{ERM}} + \xi_i.$$

Panel A reports the estimated intercept, λ^0, and the prices of risk, λ^C and λ^{ERM}. The quarterly price of risk on the orthogonal labor income return is -0.241 and significant at the 1%. This is consistent with the evidence in Gómez, Priestley, and Zapatero (2009) of an aggregate U.S. level negative price of risk associated with relative wealth concerns. The size of the estimated coefficient is larger than the state level estimate of 0.198 in Table 1, which in isolation would suggest that stocks that hedge aggregate relative wealth concerns have higher demand.

In Panel B, for each individual stock i, we estimate the rolling betas with respect to the orthogonal income return of the state where the firm headquarters are located, the orthogonal country labor income return, and the stock market return, using nine years of rolling observations (36 quarters):

$$r_{i,t} \;\; = \;\; \alpha_i + \beta_i^F r_t^F + \beta_i^C r_t^C + \beta_i^{\mathrm{ERM}} r_{\mathrm{ERM},t} + u_{i,t}.$$

We then estimate the cross sectional regression every quarter:

$$r_i \;\; = \;\; \lambda_t^0 + \lambda_t^F \hat{\beta}_i^F + \lambda_t^C \hat{\beta}_i^C + \lambda_t^{\mathrm{ERM}} \hat{\beta}_i^{\mathrm{ERM}} + \xi_i.$$

Both the country and state factors carry a negative price of risk. The size for the state factor, -0.174, is similar to the price of risk reported in Table 1. The size of the estimated price of risk on the orthogonal country level labor income relative to when it is included on its own is smaller but remains statistically significant.

As a robustness test, the state orthogonal income is replaced with the orthogonal divisional labor income. Both the orthogonal state and country prices of risk reported in Panel C decrease marginally, but remain strongly significant. These results suggest that both deviations from the Joneses consumption at the local (divisional) and country level are priced.

One way to disentangle both effects is to test the model division by division. In Table 6 Panel D, we report the cross-sectional prices of risk on the orthogonal country labor income estimated using only stocks within each division. In all divisions the price of risk is negative and strongly significant. It is worth noting that the size of these premia is very uniform across divisions, consistent with the country-wide nature of the Joneses risk considered in this test. In Panel E we observe that the risk premia vary considerably from division to division, consistent with the tests in the previous section. Five of the orthogonal state factors (MA, SA, PA, ES, WS) carry a negative premium statistically significant at least at the 5% level.

These tests corroborate the country-wide evidence in favor of the KEEPM already documented in Gómez, Priestley, and Zapatero (2009). Moreover, they show that there exists a local hedging component that varies in magnitude and power across divisions.

VII. Conclusion

Mayers (1972) pointed out the importance of human capital as a component of aggregate wealth. Following up on this idea, the finance literature has used labor income as an indicator of human capital and linked it to the cross-section of stock returns. In this paper, we show that relative wealth concerns can explain the link between labor income and stock returns.

In this paper, we show that there are local sources of relative wealth concerns that are priced in the cross section of stock returns. State level orthogonal labor income is an important determinant of the cross section of returns. In particular, the risk premium associated with labor income is negative and, even more importantly, the risk factor is local, as consistent with the economic nature of relative wealth concerns. We also document that the empirical implications of the model vary across different regions, depending on the size of the risk factor and its variability, as predicted by the model. In general, local labor income has higher correlation with local stock returns than with stock returns of other divisions, as we show in this paper. However, as we clearly document, the pricing factor is the correlation between stock returns and labor income, and not geographic location. This is clearly different from the notion of familiarity suggested in the literature as a possible factor.

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Table 1

Individual stocks

The KEEPM model: Aggregate

Let $r_{s,t}^{F}$ denote the orthogonal labor income return in state s and period t; $r_{\mathrm{RM},t}$ denotes the return on the aggregate, country stock market index. For each individual stock i we estimate the rolling betas with respect to the orthogonal labor income return from the state where the stock headquarters are located and the stock market return, using nine years of rolling observations (36 quarters). The slope coefficient $\hat{\beta}_i^F$ is estimated for every stock in the U.S. Every time a new year of quarterly data is added, the first (oldest) year is removed and the process is repeated. The time series of quarterly estimated betas starts in 1965Q1 and ends in 2011Q4. We then run cross-sectional regressions at each quarter of all U.S. individual stock returns on their estimated betas:

$$r_i \quad = \quad \lambda_t^0 + \lambda_t^F \hat{\beta}_i^F + \lambda_t^{\mathrm{RM}} \hat{\beta}_i^{\mathrm{RM}} + \xi_i.$$

The table reports the average (percentage) quarterly prices of risk λ^0, λ^F and λ^{RM}. Absolute t-values are reported below in parentheses.

Average prices of risk

λ^0	λ^F	λ^{RM}
1.993	-0.198	1.403
(4.64)	(2.27)	(1.92)

Table 2

Individual stocks

The KEEPM model: Per division

There are nine Census Bureau Divisions which we index with two capital letters: West South Central (WS), Pacific (PA), East South Central (ES), Mountain (MO), East North Central (EN), South Atlantic (SA), West North Central (WN), Middle Atlantic (MA), and New England (NE). First, stocks are sorted into divisions according to the location of the company's headquarters. We then follow the same procedure described in Table 1 and estimate the slope coefficient $\hat{\beta}_i^F$ for every stock i with respect to the orthogonal labor income return from the state where the stock headquarters are located. We then run the contemporaneous Fama-MacBeth cross-sectional regressions at each period t across stocks i in each division k:

$$r_{i,k} = \lambda_{t,k}^0 + \lambda_{t,k}^F \hat{\beta}_i^F + \lambda_{t,k}^{\mathrm{ERM}} \hat{\beta}_i^{\mathrm{ERM}} + \xi_{i,k}.$$

Panel A reports the average (percentage) quarterly prices of risk λ_k^0, λ_k^F and λ_k^{ERM} for each division k. Absolute t-values are reported in parenthesis. In Panel B we repeat the same procedure as in Panel A but, in this case, the orthogonal state labor income is replaced with the orthogonal divisional labor income. Simultaneously, the betas in each division are estimated using all stocks in the U.S., regardless of the location of their headquarters.

Average prices of risk

	Panel A			Panel B		
	Only stocks			All stocks		
	within the division			across divisions		
	State labor income			Divisional labor income		
	λ_k^0	λ_k^F	λ_k^{ERM}	λ_k^0	λ_k^F	λ_k^{ERM}
MA	1.965 (3.83)	−0.165 (1.62)	1.390 (1.80)	1.889 (3.67)	−0.187 (1.87)	1.450 (1.84)
NE	2.132 (4.71)	−0.172 (1.84)	1.473 (2.00)	2.180 (4.79)	−0.157 (1.72)	1.435 (1.95)
SA	1.988 (4.15)	−0.284 (2.71)	1.052 (1.38)	1.923 (3.98)	−0.277 (2.46)	1.203 (1.59)
EN	2.058 (4.55)	−0.178 (1.85)	1.092 (1.35)	1.942 (4.28)	−0.107 (1.20)	1.261 (1.57)
PA	1.951 (3.64)	−0.320 (2.58)	1.375 (1.76)	1.927 (3.55)	−0.315 (2.56)	1.423 (1.82)
ES	2.745 (3.28)	−0.279 (2.26)	0.542 (0.50)	2.802 (3.25)	−0.379 (2.77)	0.653 (0.58)
WS	1.675 (2.97)	−0.248 (1.87)	2.123 (2.58)	1.628 (2.85)	−0.244 (1.73)	2.154 (2.61)
WN	1.709 (3.97)	−0.209 (1.50)	1.535 (2.12)	1.755 (3.97)	−0.159 (1.23)	1.503 (2.05)
MO	2.785 (3.23)	−0.324 (2.05)	0.017 (0.02)	2.835 (3.38)	−0.343 (2.08)	0.079 (0.08)

Table 3

Beta-sorted portfolios

Time-series regressions: Per division

In each divisions, stocks are sorted according to their slope coefficients $\hat{\beta}_i^F$ estimated in equation (8) into three and ten equally weighted portfolios denoted by subscript p. The quarterly return on these portfolios is estimated over the following year. The time series of estimated quarterly returns starts in 1966Q1 and ends in 2011Q4. Panel A reports, for each division, the average percentage return on the difference between the portfolio that includes the stocks with the highest betas (P1) and the portfolio with the lowest betas (P3 and P10, respectively). Absolute t-statistics that test whether the difference between the two portfolios is different from zero are reported in parenthesis. In Panel B we repeat the same procedure as in Panel A but, in this case, the orthogonal state labor income is replaced with the orthogonal divisional labor income. Simultaneously, the betas in each division are estimated using all stocks in the U.S., regardless of the location of their headquarters. Panel C reports the average percentage return on the ten portfolios constructed with the orthogonal state income betas. Absolute t-statistics are reported in parenthesis. We next estimate, for each division, a factor mimicking (FM) portfolio for the orthogonal state labor income risk by going long on the top portfolio containing one-third of the stocks with the highest beta (P1) and short on the bottom portfolio containing one-third of the stocks with the lowest beta (P3) in the division. Let $r_{k,t}^{\mathrm{FM}}$ denote the return on the factor mimicking portfolio from division k. Panel D reports, for each division, full sample estimates of the coefficient from the following regression (absolute t-values in parenthesis):

$$r_{p,k,t} \quad = \quad \alpha_{p,k} + \beta_{p,k}^{\mathrm{FM}} r_{k,t}^{\mathrm{FM}} + \beta_{k,p}^{\mathrm{ERM}} r_{\mathrm{ERM,t}} + u_{p,k,t}.$$

Average return spread				
	Panel A		Panel B	
	Only stocks		All stocks	
	within the division		across divisions	
	State labor income		Divisional labor income	
	$P1 - P3$	$P1 - P10$	$P1 - P3$	$P1 - P10$
MA	−0.802 (1.82)	−1.045 (1.55)	−0.885 (1.79)	−0.894 (1.22)
NE	−0.429 (1.01)	−0.793 (0.93)	−0.877 (1.92)	−0.963 (1.46)
SA	−0.896 (1.60)	−1.669 (1.76)	−1.266 (2.31)	−1.833 (2.30)
EN	−0.229 (0.66)	-0.692 (1.03)	−1.041 (2.26)	−1.205 (1.78)
PA	−2.081 (3.13)	−1.857 (1.74)	−1.373 (2.66)	−1.981 (2.63)
ES	−1.335 (2.05)	−0.915 (1.24)	−0.924 (1.76)	−1.008 (1.33)
WS	−1.704 (2.59)	−2.987 (2.76)	−1.244 (2.36)	−1.658 (2.20)
WN	−0.327 (0.72)	−0.359 (0.48)	−1.090 (2.27)	−1.225 (1.73)
MO	−0.987 (1.43)	−2.523 (1.82)	−1.432 (2.64)	−1.727 (2.23)

Panel C: Portfolio returns

	MA	NE	SA	EN	PA	ES	WS	WN	MO
P1	1.963	2.591	1.751	2.058	1.951	2.386	1.413	2.062	0.526
	(2.19)	(2.69)	(2.14)	(1.91)	(1.83)	(2.80)	(1.39)	(2.39)	(0.45)
P2	2.291	2.221	2.104	2.651	2.089	2.292	1.939	2.069	2.474
	(2.96)	(2.67)	(2.81)	(2.84)	(2.35)	(3.23)	(2.26)	(2.78)	(2.21)
P3	2.162	2.852	1.878	2.622	2.599	2.376	1.734	1.615	1.989
	(2.68)	(2.92)	(2.41)	(3.08)	(2.86)	(3.37)	(2.14)	(2.43)	(2.04)
P4	2.149	2.787	2.396	2.146	2.331	2.197	2.272	2.124	1.787
	(2.58)	(3.25)	(3.11)	(2.50)	(2.62)	(2.92)	(2.82)	(3.01)	(1.95)
P5	2.053	2.560	2.218	2.301	2.665	2.383	2.711	2.313	1.852
	(2.47)	(2.80)	(2.42)	(2.87)	(2.69)	(2.95)	(3.06)	(2.78)	(1.84)
P6	2.994	2.822	2.644	2.015	2.243	2.659	2.714	2.256	1.883
	(3.48)	(2.93)	(3.02)	(2.64)	(2.29)	(3.17)	(3.24)	(2.87)	(1.69)
P7	2.458	3.415	2.945	1.772	4.118	2.569	2.701	2.621	2.237
	(2.70)	(3.46)	(3.04)	(2.21)	(3.50)	(2.80)	(3.18)	(3.26)	(2.19)
P8	2.594	2.685	2.597	1.951	3.921	2.793	3.017	2.333	3.186
	(2.68)	(2.64)	(2.54)	(2.67)	(3.44)	(2.77)	(3.24)	(2.53)	(2.45)
P9	3.612	2.997	2.649	1.700	4.515	3.249	3.236	2.270	2.685
	(3.24)	(2.70)	(2.46)	(2.30)	(3.32)	(2.93)	(2.81)	(2.25)	(2.11)
P10	3.008	3.385	3.421	2.750	3.808	3.301	4.400	2.422	3.056
	(2.32)	(2.55)	(2.58)	(2.96)	(2.46)	(2.47)	(3.39)	(2.18)	(2.22)

Panel D: $\hat{\beta}^{\mathrm{FM}}_{p,k}$

	MA	NE	SA	EN	PA	ES	WS	WN	MO
P1	−0.204	0.068	0.158	0.547	0.137	−0.019	0.580	0.102	0.404
	(2.34)	(0.63)	(2.06)	(8.76)	(1.75)	(0.32)	(7.35)	(0.96)	(4.31)
P2	−0.038	0.001	0.080	0.072	0.172	−0.074	0.442	0.241	0.304
	(0.58)	(0.02)	(1.25)	(0.81)	(2.85)	(1.72)	(7.44)	(2.76)	(3.55)
P3	−0.288	0.079	0.029	0.0644	−0.111	−0.102	0.344	0.059	−0.079
	(3.62)	(0.76)	(0.42)	(0.69)	(1.80)	(2.49)	(5.64)	(0.84)	(1.03)
P4	−0.288	−0.282	−0.145	−0.030	−0.160	−0.085	0.192	0.006	−0.025
	(3.62)	(3.08)	(2.06)	(0.30)	(2.89)	(2.10)	(3.34)	(0.08)	(0.34)
P5	−0.422	−0.336	−0.228	−0.313	−0.389	−0.139	0.084	−0.356	−0.296
	(5.76)	(3.55)	(3.02)	(3.60)	(6.19)	(3.19)	(1.25)	(4.24)	(3.55)
P6	−0.503	−0.342	−0.341	−0.188	−0.417	−0.155	−0.074	−0.400	−0.250
	(6.37)	(3.09)	(4.89)	(1.86)	(6.06)	(3.19)	(1.15)	(4.96)	(2.80)
P7	−0.595	−0.552	−0.493	−0.493	−0.601	−0.211	−0.113	−0.396	−0.313
	(7.27)	(5.18)	(6.60)	(5.08)	(8.52)	(3.84)	(1.83)	(4.43)	(3.99)
P8	−1.043	−0.768	−0.775	−0.611	−0.625	−0.267	−0.343	−0.735	−0.759
	(13.52)	(8.51)	(10.78)	(6.59)	(9.94)	(4.16)	(5.40)	(8.55)	(7.90)
P9	−1.238	−0.850	−0.851	−0.749	−1.091	−0.312	−0.578	−0.864	−0.832
	(15.36)	(8.81)	(11.17)	(7.32)	(15.62)	(4.39)	(7.57)	(9.13)	(9.19)
P10	−1.406	−1.411	−1.227	−1.021	−1.141	−0.369	−0.856	−0.950	−0.914
	(14.43)	(12.18)	(13.64)	(12.34)	(12.09)	(4.26)	(9.56)	(9.51)	(8.67)

Table 4

Beta-sorted portfolios

Cross-Sectional regressions: Per division

In Panel A, we estimate in each division the contemporaneous Fama-MacBeth cross-sectional regressions at each period t:

$$r_{p,k} \quad = \quad \lambda^{0}_{t,k} + \lambda^{\text{FM}}_{t,k}\hat{\beta}^{\text{FM}}_{p,k} + \lambda^{\text{ERM}}_{t,k}\hat{\beta}^{\text{ERM}}_{p,k} + \xi_{p,k},$$

across portfolios p in each division k. As testing portfolios we use the ten state beta-sorted portfolios from Table 3 Panel C. $\hat{\beta}^{\text{FM}}$ is the estimated beta for the factor mimicking portfolio from Table 3 Panel D; $\hat{\beta}^{\text{ERM}}$ is the estimated beta for the market risk factor. We report the average cross-sectional (percentage) quarterly prices of risks. $\overline{R}^2$ is calculated R^2 as $\left[Var_c(\overline{r}_p) - Var_c(\overline{\xi}_p)\right]/Var_c(\overline{r}_p)$, where Var_c is the cross-sectional variance, $\overline{r}_p$ is the average return and $\overline{\xi}_p$ is the average residual. $\overline{R}^2$ is the adjusted R^2. The pricing errors ($p.e.$) of a given portfolio are defined as the difference between the actual portfolio return and the expected return according to the corresponding cross-sectional model. The $p.e.$ *Test* is a Chi-sq test given as $\widehat{\alpha}' cov(\widehat{\alpha})^{-1}\widehat{\alpha}$, where $\widehat{\alpha}$ is the vector of average pricing errors across the forty-five portfolios and cov is the covariance matrix of the pricing errors. Absolute t-values in parenthesis; p-values in brackets. Panel B presents the covariances (lower triangular matrix), variances (diagonal), and correlations (upper triangular matrix) among the divisional factor mimicking portfolios defined in Table 3. In Panel C we test the KEEPM augmented with the Fama-French factors (KEEPM-FF):

$$r_{p,k} \;=\; \lambda^0_{t,k} + \lambda^{\mathrm{FM}}_{t,k}\hat{\beta}^{\mathrm{FM}}_{p,k} + \lambda^{\mathrm{ERM}}_{t,k}\hat{\beta}^{\mathrm{ERM}}_{p,k} + \lambda^{\mathrm{SMB}}_{t,k}\hat{\beta}^{\mathrm{SMB}}_{p,k} + \lambda^{\mathrm{HML}}_{t,k}\hat{\beta}^{\mathrm{HML}}_{p,k} + \xi_{p,k}.$$

Panel A. Prices of risk: KEEPM

	λ_k^0	λ_k^{FM}	λ_k^{ERM}	$\overline{R}^2$	*p.e.*	*p.e. Test*
MA	2.406	−0.903	−0.412	0.57	0.261	13.120
	(1.51)	(2.01)	(0.24)			[0.11]
NE	2.055	−0.500	0.463	0.40	0.190	3.664
	(1.03)	(1.11)	(0.24)			[0.89]
SA	2.059	−0.991	0.025	0.77	0.196	3.427
	(1.03)	(1.73)	(0.01)			[0.90]
EN	−0.216	−0.101	2.291	0.17	0.253	11.117
	(0.13)	(0.28)	(1.28)			[0.20]
PA	0.449	−1.812	1.545	0.76	0.366	0.079
	(0.21)	(2.63)	(0.75)			[1.00]
ES	1.117	−1.456	1.046	0.89	0.103	7.266
	(0.88)	(0.60)	(0.64)			[0.51]
WS	2.795	−1.805	−0.140	0.92	0.174	2.219
	(1.31)	(2.68)	(0.07)			[0.97]
WN	2.366	−0.413	−0.316	0.34	0.132	2.766
	(1.26)	(0.89)	(0.14)			[0.95]
MO	−0.546	−1.193	2.199	0.63	0.302	2.979
	(0.19)	(1.77)	(0.79)			[0.94]

Panel B. Variances, Covariances and Correlations

	MA	NE	SA	EN	PA	ES	WS	WN	MO
MA	0.0037	0.4350	0.6546	0.4264	0.4546	0.3731	0.5446	0.5543	0.3088
NE	0.0015	0.0033	0.4059	0.3648	0.4670	0.1965	0.4785	0.3655	0.2474
SA	0.0030	0.0018	0.0059	0.5041	0.5047	0.2699	0.5840	0.5214	0.2533
EN	0.0012	0.0009	0.0018	0.0022	0.5040	0.3422	0.4388	0.4810	0.2584
PA	0.0025	0.0025	0.0036	0.0022	0.0086	0.2480	0.6460	0.5496	0.1620
ES	0.0020	0.0010	0.0018	0.0014	0.0021	0.0083	0.3369	0.2663	0.1353
WS	0.0030	0.0025	0.0041	0.0018	0.0055	0.0028	0.0083	0.5729	0.2675
WN	0.0020	0.0013	0.0024	0.0013	0.0031	0.0015	0.0032	0.0038	0.2286
MO	0.0017	0.0013	0.0018	0.0011	0.0014	0.0011	0.0023	0.0013	0.0089

Panel C. Prices of risk: KEEPM-FF

	λ_k^0	λ_k^{FM}	λ_k^{ERM}	λ_k^{SMB}	λ_k^{HML}	$\overline{R}^2$	*p.e.*	*p.e. Test*
MA	1.455	−0.938	1.632	−0.969	−0.139	0.64	0.241	14.228
	(0.71)	(2.08)	(0.57)	(0.52)	(0.07)			[0.03]
NE	0.693	−0.519	0.098	1.869	1.219	0.78	0.139	1.861
	(0.32)	(1.16)	(0.04)	(1.35)	(0.87)			[0.93]
SA	0.907	−1.009	0.008	1.193	1.228	0.80	0.176	2.482
	(0.36)	(1.77)	(0.00)	(0.72)	(0.79)			[0.87]
EN	−2.883	−0.169	7.678	−1.679	−3.403	0.48	0.232	7.013
	(1.06)	(0.48)	(2.31)	(1.25)	(1.64)			[0.32]
PA	1.307	−1.807	2.308	−0.902	−1.293	0.81	0.320	8.448
	(0.49)	(2.61)	(0.83)	(0.51)	(0.67)			[0.21]
ES	2.671	−2.771	−0.352	0.571	−0.721	0.88	0.085	5.771
	(1.55)	(1.20)	(0.21)	(0.45)	(0.43)			[0.45]
WS	3.696	−1.797	−0.588	0.322	−1.017	0.92	0.152	2.201
	(1.39)	(2.69)	(0.28)	(0.30)	(0.39)			[0.90]
WN	1.755	−0.423	−0.582	0.950	1.513	0.61	0.117	1.809
	(0.64)	(0.91)	(0.22)	(0.87)	(0.69)			[0.94]
MO	−0.714	−1.232	2.044	1.275	−0.436	0.64	0.299	2.795
	(0.22)	(1.66)	(0.69)	(0.77)	(0.28)			[0.83]

Table 5

GMM estimation of the Joneses parameter

We derive the following system system of orthogonality conditions from the equilibrium condition of equations (4) and (5):

$$
\begin{aligned}
0 &= \lambda^{\mathrm{ERM}} - H\left(1 - b\sum_{k=1}^{K}\omega_k\theta_k\,\beta_k\right)\left(r_M - \mathrm{E}(r_M)\right)^2,\\
0 &= \lambda_k^{\mathrm{FM}} + H\,b\left(\omega_k\theta_k\,\left(r_k^{\mathrm{FM}} - \mathrm{E}(r_k^{\mathrm{FM}})\right)^2 + \sum_{k'\neq k}\omega_{k'}\theta_{k'}\,\left(r_k^{\mathrm{FM}} - \mathrm{E}(r_k^{\mathrm{FM}})\right)\left(r_{k'}^{\mathrm{FM}} - \mathrm{E}(r_{k'}^{\mathrm{FM}})\right)\right),
\end{aligned}
$$

for each division $k = 1, 2, ..., K$. $b = \frac{\gamma}{1-\gamma}$. r_M and r_k^{FM} denote, respectively, the time series of the U.S. market return and the return on the factor mimicking portfolio from division k, from the first quarter of 1965 to the final quarter of 2011. We take λ_k^{FM} from Table 4 Panel A. θ_k is proxied using the time series of divisional personal income as a proportion of the divisional GDP. To proxy ω_k, each quarter, the market capitalization of all stocks in the division is divided by the aggregate market capitalization. We assume different values of the aggregate risk aversion coefficient H and estimate the parameter γ and the market price of risk, λ^{ERM} using Hansen's Generalized Method of Moments (GMM). Hansen's J-test tests whether the orthogonality conditions are jointly zero. It follows a $\chi^2(N - L)$ distribution. Standard errors in parenthesis; p-values in brackets.

$H=1$		$H=3$		$H=6$		J-test
λ^{ERM}	γ	λ^{ERM}	γ	λ^{ERM}	γ	
-0.0573	0.996	-0.0421	0.989	-0.0194	0.978	183.9
(0.0103)	(0.000488)	(0.00940)	(0.00144)	(0.00864)	(0.00282)	[0.000]

Table 6

Individual stocks

Country-wide orthogonal income

Let r_t^C denote the orthogonal country labor income return in the U.S. in period t; $r_{\mathrm{ERM},t}$ denotes the return on the aggregate, country stock market index. For each individual stock i we estimate the rolling betas with respect to the orthogonal country labor income return and the stock market return, using the same procedure as in Table 1:

$$r_{i,t} \;=\; \alpha_i + \beta_i^C r_t^C + \beta_i^{\mathrm{ERM}} r_{\mathrm{ERM},t} + u_{i,t}.$$

The slope coefficients $\hat{\beta}_i^C$ and $\hat{\beta}_i^{\mathrm{ERM}}$ are estimated for every stock in the U.S. The time series of quarterly estimated betas starts in 1965Q1 and ends in 2011Q4. We then run the contemporaneous Fama-MacBeth cross-sectional regressions at each period t across stocks i:

$$r_i \;=\; \lambda_t^0 + \lambda_t^C \hat{\beta}_i^C + \lambda_t^{\mathrm{ERM}} \hat{\beta}_i^{\mathrm{ERM}} + \xi_i.$$

Panel A reports the average (percentage) quarterly prices of risk λ^0, λ^C and

λ^{ERM}. Absolute t-values are reported in parenthesis. In Panel B, for each individual stock i we estimate the rolling betas with respect to the orthogonal income return of the state where the headquarters are located, the orthogonal country labor income return and the stock market return, using nine years of rolling observations (36 quarters):

$$r_{i,t} = \alpha_i + \beta_i^F r_t^F + \beta_i^C r_t^C + \beta_i^{\mathrm{ERM}} r_{\mathrm{ERM},t} + u_{i,t}.$$

In Panel C the state orthogonal income is replaced with the orthogonal divisional labor income. Each panel reports the corresponding cross-sectional (percentage) prices of risk. Panels D and E repeat the analysis within each division. In Panel D, only orthogonal country income is considered. In Panel E we include the orthogonal state risk factor in each division. In both cases, only stocks within each division are included.

Country-wide orthogonal income		
Panel A. All U.S. stocks		
λ^0	λ^C	λ^{ERM}
1.797 (3.88)	-0.241 (2.56)	1.335 (1.82)

Local and country-wide orthogonal income

Panel B State orthogonal labor income				Panel C Divisional orthogonal labor income			
λ^0	λ^F	λ^C	λ^{ERM}	λ^0	λ^F	λ^C	λ^{ERM}
1.952 (4.76)	-0.174 (2.24)	-0.192 (2.36)	1.317 (1.83)	2.019 (4.95)	-0.161 (2.07)	-0.170 (2.10)	1.295 (1.81)

Divisional tests

	Panel D Only country-wide orthogonal income			Panel E Country-wide and state orthogonal income			
	λ_k^0	λ_k^C	λ_k^{ERM}	λ_k^0	λ_k^F	λ_k^C	λ_k^{ERM}
MA	1.202 (1.96)	−0.251 (2.28)	1.771 (2.18)	2.087 (4.49)	−0.69 (1.72)	−0.176 (1.87)	1.196 (1.87)
NE	2.609 (5.04)	−0.182 (1.89)	0.862 (1.15)	2.076 (4.61)	−0.136 (1.52)	−0.196 (2.10)	1.234 (1.71)
SA	1.940 (3.75)	−0.214 (2.28)	1.106 (1.48)	1.743 (3.75)	−0.282 (2.72)	−0.229 (2.42)	1.230 (1.65)
EN	1.958 (4.10)	−0.206 (2.09)	1.054 (1.29)	2.285 (5.29)	−0.131 (1.47)	−0.171 (1.92)	0.748 (0.95)
PA	1.674 (2.66)	−0.265 (2.29)	1.492 (1.84)	2.184 (4.10)	−0.318 (2.60)	−0.238 (2.37)	1.158 (1.51)
ES	2.302 (3.11)	−0.234 (2.08)	1.207 (1.21)	4.031 (3.01)	−0.406 (2.35)	−0.219 (1.92)	−0.819 (0.50)
WS	1.611 (2.71)	−0.244 (2.46)	2.036 (2.46)	1.557 (2.99)	−0.235 (1.75)	−0.241 (2.58)	2.015 (2.48)
WN	1.992 (4.39)	−0.206 (1.89)	1.179 (1.64)	2.012 (4.80)	−0.177 (1.35)	−0.227 (2.20)	1.076 (1.52)
MO	3.195 (3.53)	−0.229 (2.12)	−0.194 (0.19)	2.968 (3.97)	−0.186 (1.21)	−0.133 (1.24)	0.134 (0.14)

Internet Appendix: Labor Income, Relative Wealth Concerns, and the Cross-section of Stock Returns

This Appendix includes some results that, although not essential to understand the main arguments of the paper, provides complementary material and further support to the evidence presented in the body of the paper. The Appendix has three main parts, and two figures referenced in the paper. In Appendix A, we summarize the optimal portfolio choice problem of an agent with either endogenous or exogenous keeping up with the Joneses preferences. For a detailed derivation, we refer the reader to Gómez, Priestley and Zapatero (2009). In Appendix B we present the portfolio tests of the country level analysis. This is a robustness test. We also perform portfolio tests at the divisional level, but these are less likely to be biased by the sorting procedure, therefore we leave them in the body of the paper. In Appendix C we derive a linear approximation to the stochastic discount factor (SDF) implied by our model. We then apply the SDF to the estimation of the prices of risk by the General Method of Moments (GMM).

Appendix A

0.1 Exogenous keeping up with the Joneses preferences

In this subsection we analyze the implications of a version of the keeping up with the Joneses preferences of Abel (1990) and Galí (1994). In particular, in the economy we consider investors are endowed with an utility function[1]

$$u(c, C) = \frac{c^{(1-\alpha)}}{1-\alpha} C^{\gamma\alpha}, \tag{1}$$

where c denotes the investor's consumption of the single consumption good, the economy's numeraire; C is the division average or per capita consumption; $\alpha > 0$ is the (constant) relative risk-aversion coefficient and $1 > \gamma \geq 0$ is the "Joneses parameter."

Here, workers represent agents endowed with non-tradable income. For instance, their human capital, that will materialize into wage income, or entrepreneurial income. Call w_k^0 the initial aggregate endowment of non-financial wealth for workers in division k; w_k denotes the final ($t = 1$) random value of their non-tradable income. Workers face incomplete markets because they cannot trade their human capital (due to moral hazard issues) and have no access to financial markets; therefore, they cannot hedge their income risk.

Since each investor takes C as exogenous and common, the typical aggregation property of the CRRA utility functions allows us to replace all the investors in a given division by a representative investor with utility function (1) endowed with the aggregated investors income without affecting the equilibrium prices. At time $t = 0$ each representative investor is endowed with a share of the local firm (unit value by assumption); hence, $c^0 = 1$ in all divisions.

We can write the problem's first order condition as a function of the investor's consumption and the workers relative wealth, w/c:

$$E\left(r\, c^{-\alpha(1-\gamma)}(1 + w/c)^{\alpha\gamma}\right) = 0. \tag{2}$$

[1] To simplify the notation, we drop the division subindex k for the moment (thus, all variables to be introduced next apply to investors in any division).

Notice that, in the absence of keeping up with the Joneses behavior ($\gamma = 0$), the previous condition reduces to $E\left(r\, c^{-\alpha}\right) = 0$, the standard CRRA Euler equation.

Condition (2) allows us to solve for the representative investor's optimal portfolio. Since financial markets are complete, there exists a mimicking portfolio X^w that maps the workers relative income onto the investment opportunity set such that $w/c = w^0(R + r'X^w)$. Following Galí (1994), given w^0 and X^w, for small values of $E(r)$, the optimal portfolio of the representative investor of division k can be approximated as a function of α, γ and the risk adjusted risk premia $\Omega^{-1}E(r)$:

(3) $$x_k^* = \frac{\theta_k \gamma_k}{1-\gamma_k} X_k^w + \frac{1}{\alpha_k(1-\gamma_k)} \Omega^{-1} E(r),$$

with $\theta_k = \frac{w_k^0}{1+w_k^0}$, the workers initial wealth as a proportion of the division's total wealth (investor's plus non-diversifiable wealth).

Notice that even if there is a friction ($\theta_k > 0$) that prevents full risk-diversification for a set of agents (the workers), investors will hold well diversified portfolios unless they exhibit some degree of keeping up with the Joneses behavior ($\gamma_k > 0$). Thus, it is important to emphasize that investors' portfolios will be locally biased if and only if *both* keeping up with the Joneses behavior and a market friction exist.

0.2 Endogenous keeping up with the Joneses preferences

In this section, we discuss the endogenous keeping up with the Joneses preferences presented in DeMarzo, Kaniel and Kremer (2004). In this specification agents consume two types of goods: c, which has the interpretation of a global good, and w_k, a local good, like housing services. Utility over consumption for these two goods is given by:

$$u(c, w) = \frac{1}{1-\alpha}(c^{1-\alpha} + \delta w^{1-\alpha}).$$

The parameter $\delta > 0$ specifies the relative importance of the local good. All consumption takes place at the end of the period. At time $t = 0$, investors are endowed with shares of the firm that produces the global good. Call c_k^0 the aggregate value of those shares at the beginning of the period for agents in division k. For simplicity, let $c_k^0 = 1$ in all divisions.

Workers in each division will receive a fixed number $\bar{w}_k$ of units of the local good at time $t = 1$. In equilibrium, the relative price of the local good in terms of the global good at $t = 1$ is given by $p_k = \delta \left(\frac{c_k}{\bar{w}_k}\right)^{\alpha}$. As it would be expected, the scarcer the (fixed) local good endowment relative to the (stochastic) global good consumption, the higher the relative price of the former. The investor's hedging demand for this risk will trigger the endogenous keeping up with the Joneses behavior in this model. Financial markets are complete.

If workers can not diversify their endowment risk (due, for instance, to short-selling constraints and moral hazard), Proposition 2 in DeMarzo, Kaniel and Kremer (2004) shows that the representative investor's marginal utility is given by:

$$u_c(c, p) = c^{-\alpha} \left(1 + \delta^{1/\alpha} p^{1-1/\alpha}\right)^{\alpha}. \tag{4}$$

Let $p^0 = \delta \left(\frac{c^0}{\bar{w}}\right)^{\alpha}$ denote the relative price at $t = 0$ of one unit of the non-diversifiable, local good endowment of workers at time $t = 1$. Recall that we normalized the initial investor's shares endowment $c^0 = 1$. Hence, $p^0 = \delta \bar{w}^{-\alpha}$. The present value of the workers endowment is therefore $\bar{w}^0 = \delta \bar{w}^{1-\alpha}$.

In this model, the relative wealth at $t = 0$ of the workers in division k as a proportion of the total division wealth is given by $\theta_k = \frac{\bar{w}_k^0}{1+\bar{w}_k^0}$. Call $\bar{w}_k p_k / \bar{w}_k^0$ the return on the workers wealth (in units of the global good) over the period. Like in the exogenous preferences specification, under complete (financial) markets, there exists a portfolio X_k^w such that $\frac{\bar{w}_k p_k}{\bar{w}_k^0} = R + r' X_k^w$.

After these definitions, we can write the approximate function for division k investor's optimal portfolio as follows:

$$x_k^* = \frac{\theta_k(\alpha_k - 1)}{\alpha_k} X_k^w + \frac{1}{\alpha_k} \Omega^{-1} E(r). \tag{5}$$

Notice that, in this model, the optimal portfolio for the logarithmic investor ($\alpha = 1$) coincides with the benchmark, well diversified portfolio $\Omega^{-1}E(r)$. No relative wealth concern arises even in the presence of local, non-diversifiable wealth. Only for $\alpha > 1$ should we observe a local bias in portfolio holdings.

Appendix B

0.1 Country Level Portfolios

We complement here the results of Section 4 in the paper. In that section we test the implications of the model using individual stocks. In addition, it is standard in empirical asset pricing tests to use portfolios of stocks as test assets in order to reduce the errors in variables problem that plagues the two-step Fama and MacBeth (1973) estimation methodology. Furthermore, it is common to use factor mimicking portfolios to proxy risk factors in order to be able to interpret the estimated prices of risk in terms of returns (risk premia). Furthermore, model performance that focuses on pricing errors is easier to undertake with the use of well diversified portfolios.

We construct a factor mimicking (FM) portfolio for the orthogonal state labor income risk as follows. For each stock i, we use the slope coefficient on the orthogonal labor factor $\hat{\beta}_i^F$ estimated in equation (6) in the paper until the fourth quarter of 1964, to rank stocks in 1965. Next, we form three equally weighted portfolios according to the size of the coefficient. We then add one year of quarterly data. We re-estimate the coefficient, rank the stocks, sort them into three portfolios and compute their quarterly returns in 1966. We continue adding one year and re-estimating the coefficients until we have thirty-six quarterly observations in the time-series regressions. At this point, we start rolling the data one year at a time: adding on a new year and taking off the first year. We continue this process until the end of the sample.

The above procedure results in three portfolios, from the first quarter of 1965 to the final quarter of 2011, formed in year t based on the estimated coefficient on orthogonal labor income estimated until year $t-1$. The returns of the factor mimicking portfolio are computed as the returns of the portfolio (P1) formed by the stocks with the highest one third of coefficient estimates minus the returns on the portfolio (P3) formed by the stocks with the lowest one third of coefficient estimates. We represent by r_t^{FM} the return of the state factor mimicking portfolio at t.

As test assets, we consider the Fama and French twenty-five size and book to market portfolios that have become standard in asset pricing tests due to their large spread in returns. In addition, we form test portfolios based on the sorted orthogonal betas from individual stocks estimated in equation (6) in the paper. The reason for this is that Daniel and Titman (2011) and Lewellen, Nagel, and Shanken (2010) note that testing asset pricing models using portfolios formed on firm characteristics, such as size and book to market, can lead to spurious conclusions about the usefulness of a proposed factor. This is because the factor structure of the portfolios is so strong that any proposed factor that is only weakly correlated with size or book-to-market will appear to price the test assets. That is, testing a new proposed factor on test assets sorted only by size and book-to-market is likely to have very low power. In order to alleviate this concern, we follow the recommendations in Daniel and Titman (2011) and Lewellen, Nagel, and Shanken (2010) and sort stocks by lagged loadings on our proposed factor. We use these beta-sorted portfolios together with the twenty-five Fama and French portfolios sorted by size and book-to-market in the cross-sectional tests of our model.

To generate the beta-sorted test portfolios we repeat the procedure discussed above and construct ten and twenty equally weighted portfolios.[2] We calculate excess returns on all the test portfolios by subtracting the one month T-bill rate from the actual returns.

Panel A of Table 1 in this Appendix shows the average return spread between the portfolio containing the stocks with the highest orthogonal betas (P1) and the portfolio containing the stocks with the lowest orthogonal betas (P3, P10 and P20, respectively). Notice that, consistent with the model's prediction, portfolios with a higher orthogonal beta carry a lower return relative to portfolios with lower orthogonal beta. This difference is economically significant and above 1% per quarter. We test whether the difference between both portfolios is different from zero. In the first two cases (P1-P3 and P1-P10) we strongly reject that the difference is zero. In the third case (P1-P20) we can only reject it marginally, however it should be noted that the size of the spread on the P1-P20 portfolio is larger than the spread on the P1-P3 portfolio. The lower level of statistical significance could be due to the smaller number of stocks in the 20 portfolios.

Panel B of Table 1 presents the average excess return on each of the twenty beta-sorted portfolios and the correlation coefficient between each portfolio and the factor mimicking portfolio.[3] Notice that as we move from top to bottom in the table, the average return on the portfolios increases while the correlation decreases. That is, portfolios more correlated with the factor mimicking portfolio offer a better hedging against deviations from the Joneses consumption (including non-diversifiable wealth) and trade at a higher price (lower expected return). Using the full sample, the coefficient $\hat{\beta}_p^{FM}$ is obtained by regressing the return on each of the portfolios against the return on the factor mimicking portfolio and the market excess return:

$$r_{p,t} = \alpha_p + \beta_p^{FM} r_t^{FM} + \beta_p^{erm} r_{erm,t} + u_{p,t}. \tag{1}$$

[2]All the results presented in the paper are generally robust to the use of market capitalization weighted portfolios.

[3]When examining portfolios, we use excess returns in order to test whether the models' pricing errors are equal to zero.

where $r_{p,t}$ is the excess return on portfolio p, r_t^{FM} is the return on the factor mimicking portfolio, $r_{erm,t}$ is the excess return on the aggregate stock market portfolio, and $u_{p,t}$ is the residual. All but the three top betas are strongly statistically significant. The spread in returns and betas indicates that orthogonal local labor income is closely related to stock returns. It is worth mentioning that if the distributions of betas and the prices of risk are different across divisions, the estimates of α_p from regression (1) cannot be interpreted as KEEPM pricing errors even if the KEEPM model is true. The fact that the spread in average returns in panel B of Table 1 is not monotonic suggests so.[4] Furthermore, nine of the estimate of α_p are statistically significant although the patterns is the alpha have no clear relation to the portfolios sorting. In order to address the concern that the distribution of betas and prices of risk could be different across divisions, we will repeat the time series and cross-section tests at the divisional level in section 5 in the paper.

We now turn to analyzing the cross-sectional performance of the KEEPM. As test assets, we use the excess returns on the twenty beta-sorted portfolios plus the Fama and French twenty-five size and book to market portfolios.[5] The cross sectional regressions regress excess returns in each quarter on the portfolio betas on the factor mimicking portfolio based on the orthogonal state labor income return, $\hat{\beta}_p^{FM}$, and the stock market return, $\hat{\beta}_p^{erm}$, estimated in (1):

$$r_p = \lambda_t^0 + \lambda_t^{FM}\hat{\beta}_p^{FM} + \lambda_t^{erm}\hat{\beta}_p^{erm} + \xi_p. \tag{2}$$

The results from this cross-sectional regression are reported in Panel A of Table 2 in this Appendix. The quarterly price of risk on the orthogonal factor mimicking portfolio is negative, economically important at -0.88% per quarter, and statistically significant. The adjusted R^2, $\overline{R}^2$, is 66% indicating a good measure of fit.[6] Notice that the intercept, which should be zero, is positive and statistically significant, which indicates that the model is not correctly specified. As noted above, one potential reason for this that we explore later is that we restrict the price of risk on the orthogonal labor income risk to be the same across all stocks irrespective of where the stocks come from.

[4] We thank the referee for this comment.

[5] All cross-sectional results are qualitatively analogous when the prices of risk are estimated with respect to the one-year lagged betas.

[6] Following Jagannathan and Wang (1996) and Lettau and Ludvigson (2001b), we calculate R^2 as $\left[Var_c(\overline{r}_p) - Var_c(\overline{\xi}_p)\right]/Var_c(\overline{r}_p)$, where Var_c is the cross-sectional variance, $\overline{r}_p$ is the average return and $\overline{\xi}_p$ is the average residual. $\overline{R}^2$ is the adjusted R^2.

For comparison, the second row of Panel A reports the price of risk, λ^{erm}, and the average $\overline{R}^2$ from the cross-sectional regression of the CAPM. As in previous tests of the unconditional CAPM, the estimated price of risk is statistically not different from zero and the $\overline{R}^2$ is only 8%. The third row of Panel A presents the results from the three-factor Fama and French model. The prices of risk associated with size, λ^{smb}, and book-to-market, λ^{hml}, are positive and statistically different from zero. The average $\overline{R}^2$ is 73% and similar to the value from the KEEPM.

In order to assess whether there is any additional explanatory power in the size and book-to market risk factors relative to that of the KEEPM, in the fourth row of Panel A, we estimate the KEEPM whilst including the *smb* and *hml* risk factors in (2):

$$r_p = \lambda_t^0 + \lambda_t^{FM}\hat{\beta}_p^{FM} + \lambda_t^{erm}\hat{\beta}_p^{erm} + \lambda_t^{smb}\hat{\beta}_p^{smb} + \lambda_t^{hml}\hat{\beta}_p^{hml} + \xi_p. \tag{3}$$

The price of risk on the orthogonal factor mimicking portfolio is -0.896, virtually identical to the original value reported in the first row, and also statistically significant. The estimated price of risk on the book-to-market factor remains positive and significant, although the size risk premium is smaller and only marginally significant and the $\overline{R}^2$ is 75%. We interpret these findings as evidence that the model's prediction of a negative price of risk on the Joneses risk-hedging factor remains robust to the inclusion of other risk factors known for their ability to explain the cross-section of the US stock returns. In light of the results in the fourth row of Panel A, we conclude that the orthogonal labor income factor commands a price of risk not explained by the size and book-to-market risk premia. Appendix C shows that these results are robust, and even stronger, when estimated by GMM.

To provide a more formal test of the performance of the KEEPM relative to the CAPM and the three-factor Fama and French model, Panel B presents the square root of the squared pricing errors for each test portfolio and each model. We define the pricing error of a given portfolio as the difference between the actual portfolio return and the expected return according to the cross-sectional model. Overall, the size of the pricing errors of the KEEPM are small relative to the portfolio returns in Panel B of Table 1. In particular, the average pricing error is 0.287, about ten times smaller than the average portfolio return. The pricing errors from the CAPM, as expected, are large relative to those of the other models, with an average value of 0.483. A comparison of the pricing errors of the KEEPM with those of the three-factor Fama and French model reveals that they are of similar magnitude (the average value is 0.241) and smaller in eight out of the twenty-five Fama and French portfolios (Panel

B.1) and eleven out of the twenty beta-sorted portfolios (Panel B.2). When we add the size and book-to-market risk factors to the KEEPM in the last column, the average pricing error decreases to 0.222 and the pricing errors are smaller than those from the three-factor model for eleven of the Fama and French portfolios and thirteen of the beta-sorted portfolios.

Panel B.3 includes the average cross-sectional pricing errors from each model for the forty-five portfolios of panels B.1 and B.2. We also test whether the pricing errors are jointly zero.[7] Except for the three-factor Fama and French model, the test rejects the hypothesis that the pricing errors are jointly zero, although it should be emphasized that the pricing errors are economically small and very similar in size for the KEEPM and the Fama-French three factor model and actually smaller for the model that incorporates all four factors.

The evidence presented so far shows strong support at the country level for the main prediction of the model: a negative and significant price of risk on the orthogonal state labor income return factor. In the time-series, the test portfolios' betas with respect to the orthogonal factor mimicking portfolio are strongly significant in most cases and give a reasonable and statistically significant spread in returns. In the cross-sectional tests, the KEEPM performs well both in its own right and in comparison with the three-factor Fama and French model, and the orthogonal factor mimicking portfolio is shown to be robust to the inclusion of the size and book-to-market risk factors.

[7]This is a Chi-sq test given as $\widehat{\alpha}' cov(\widehat{\alpha})^{-1}\widehat{\alpha}$, where $\widehat{\alpha}$ is the vector of average pricing errors across the forty-five portfolios and cov is the covariance matrix of the pricing errors. p-values in brackets.

Appendix C

0.1 A linear approximation to the Stochastic Discount Factor (SDF)

We will derive this approximation both for the endogenous and the exogenous case. Like in the derivation of the optimal portfolio, the only difference will lay on the interpretation of the deep parameters.

Let us start with the exogenous Joneses specification. The investor's (first order condition) optimal consumption choice is given in equation (2). This condition holds for any asset i and any division k. The first-order approximation to the marginal's utility is given by $u_c(c_k, w_k/c_k) \approx u_c(c_k^0, w_k^0/c_k^0) + u_{c,c}(c_k^0, w_k^0/c_k^0)(c_k - c_k^0) + u_{c,w/c}(c_k^0, w_k^0/c_k^0)(w_k/c_k - w_k^0/c_k^0) = u_c(1, w_k^0)\left[1 - \alpha(1-\gamma)(r'x_k^* + R - 1) + \theta_k\, \alpha\gamma(r'X_k^w + R - 1)\right]$.

Replacing the later expression in (2) we obtain the following condition:

$$E\left(r_i\ \left[\tau_k - (r'x_k^* + R - 1) + \theta_k b_k (r'X_k^w + R - 1)\right]\right) = 0, \tag{1}$$

where $\tau_k = \frac{1}{\alpha_k(1-\gamma_k)}$ and $b_k = \frac{\gamma_k}{1-\gamma_k}$. We multiply (1) by ω_k, the proportion of country market capitalization in division k and add up across all divisions:

$$E\left(r_i\left[H^{-1} - (r_M + R - 1) + \sum_k \omega_k\theta_k b_k(r_k^w + R - 1)\right]\right) = 0, \tag{2}$$

where $H^{-1} = \sum_k \omega_k\tau_k$ is the aggregate risk aversion coefficient. We have used the market clearing condition $\sum_k \omega_k x_k^* = x_M$ and the definitions $r_M = r'x_M$ and $r_k^w = r'X_k^w$. After regressing the workers non-diversifiable income onto the country market portfolio return – equation (2) in the paper– we can write $r_k^w = \beta_k r_M + r_k^F$. We replace the later expression in the Euler equation. Moreover, we assume that $E(r_i)(R-1) \approx 0$ for small values of $E(r_i)$ and the (net) risk-free rate, $R-1$. This results into the following Euler equation:

$$E\left(r_{i,t}\left[H^{-1}-(1-\sum_{k=1}^{K}\omega_k\theta_k b_k\beta_k)r_M+\sum_{k=1}^{K}\omega_k\theta_k b_k r_k^F\right]\right)=0. \tag{3}$$

We turn now to the endogenous Joneses specification. The investor's optimal consumption choice is given in equation (4). This expression can be linearly approximated as follows: $u_c(c_k,p_k)\approx u_c(c_k^0,p_k^0)+u_{c,c}(c_k^0,p_k^0)(c_k-c_k^0)+u_{c,p}(c_k^0,p_k^0)(p_k-p_k^0)$.

Replacing the values of c_k^0, p_k^0, and c_k in the later expression we obtain the following:

$$u_c(c_k,p_k)\approx u_c(1,\delta_k\bar{w}_k^{-\alpha_k})\left[1-\alpha_k(r'x_k^*+R-1)+\theta_k\left(\alpha_k-1\right)\left(\frac{\bar{w}_k p_k}{\bar{w}_k^0}-1\right)\right].$$

We replace the later expression in (4). Given that $\frac{\bar{w}_k p_k}{\bar{w}_k^0}=R+r_k'^w$, we obtain condition (1) with $\tau_k=\frac{1}{\alpha_k}$ and $b_k=\frac{\alpha_k-1}{\alpha_k}$. Following the same procedure as in the exogenous case we arrive at equation (2) and finally equation (3).

Equation (3) implies an approximate linear expression to the stochastic discount factor:

$$m\approx c_0+c_M r_M+\sum_{k=1}^{K}c_k r_k^F, \tag{4}$$

where $c_0=H^{-1}$, $c_M=\sum_{k=1}^{K}\omega_k\theta_k b_k\beta_k-1$ and $c_k=\omega_k\theta_k b_k$. We define $\boldsymbol{c}\equiv(c_M,c_1,...,c_k,...,c_K)$ as the vector of coefficients; recall that $\boldsymbol{r^F}=(r_M,r_1^F,...,r_k^F,...,r_K^F)$ denotes the vector of factor returns. Hence, we can write the stochastic discount factor is a more compact way as $m\approx c_0+\boldsymbol{c}\boldsymbol{r^F}$.

Define the forecast error at time t for the parameter vector $\boldsymbol{c}$ as $\mathbf{v}_t(\boldsymbol{c})\equiv r_t(c_0+\boldsymbol{c}\boldsymbol{r^F})$, such that, according to the equilibrium orthogonality condition (3), its unconditional expectation $E(\mathbf{v}_t(\boldsymbol{c}))=0$. Define the sample mean of the forecast errors over the T observations as:

$$\mu_T(\boldsymbol{c})\equiv\frac{1}{T}\sum_{t=1}^{T}\mathbf{v}_t(\boldsymbol{c}).$$

The GMM methodology estimates the parameter vector $\boldsymbol{c}$ that minimizes

$$\widehat{\boldsymbol{c}} = \arg\min_{c}(\mu_T'(\boldsymbol{c})\boldsymbol{\Sigma}\mu_T(\boldsymbol{c})),$$

where $\boldsymbol{\Sigma}$ is a positive definite weighting matrix. Under GMM this weighting matrix is the inverse of a consistent estimator of the spectral density matrix of $\mathbf{v}_t$ at frequency zero, defined as $\mathbf{S} = \sum_{j=-\infty}^{\infty} E[\mathbf{v}_t\mathbf{v}_{t-j}'] = T \cdot \text{var}(\mu_T)$. Hansen (1982) shows that it is optimal to use the inverse of a consistent estimator of $\mathbf{S}$ as the weighting matrix, since the estimated parameter vector has the lowest variance asymptotically. When $\boldsymbol{\Sigma}$ is the optimal weighting matrix $\mathbf{S}^{-1}$, the asymptotic standard errors are given by $\text{var}(\widehat{\boldsymbol{c}}) = \frac{1}{T}(\mathbf{D}_T'\mathbf{S}^{-1}\mathbf{D}_T)^{-1}$, where $\mathbf{D}_T = \frac{\partial\mu_T(\boldsymbol{C})}{\partial\boldsymbol{C}'}$.

0.2 GMM Estimation of the Prices of Risk

In this subsection, we approach the estimation of prices of risk implied by the KEEPM model in sections 0.1 in this Appendix and 5.2 in the paper using the GMM procedure described above. This allows for any heteroscdasticity in the residuals to be controlled for, a potential weakness of the OLS based Fama-MacBeth methodology. Whilst it is possible to use GLS in the Fama-MacBeth methodology, there are well know problems associated with the GLS method in cross sectional regressions. In particular, in finite samples the covariance matrix could be poorly estimated (see Cochrane (2005)). In order to capture possible heteroscedasticity in the residuals, we have decided to estimate the prices of risk using one-step GMM which provides robust standard errors. This approach is based on the derivation of an approximate linear stochastic discount factor (SDF) in Appendix C.1 based on the model equilibrium condition for the prices of risk (4).

Equation (2) in section 0.1 shows the Fama-MacBeth cross-section regression that we used to estimate the orthogonal price of risk on the factor mimicking portfolio (common across all portfolios), λ^{FM}, and the market price of risk, λ^{erm}. We now use the approximate Euler equation (3) and the implied stochastic discount factor, m, in (4) to estimate the corresponding loadings and prices of risk using the GMM approach. In particular, for each portfolio p, the orthogonality condition is given by:

$$r_p\ \left(c_0 + c_{erm}r_{erm} + c_{FM}r^{FM}\right) = 0, \tag{5}$$

where r^{FM} represents the return on the (orthogonal) labor risk factor-mimicking portfolio, common across all divisions; c_{FM} is the corresponding factor loading in the SDF. We also follow Cochrane (2005), who notes that, given a stochastic discount factor m, a risk factor vector $\boldsymbol{r^F}$, and factor risk premia vector $\boldsymbol{\lambda}$, $E(m\boldsymbol{r^F} - \boldsymbol{r^F} + \boldsymbol{\lambda}) = 0$. This results into two additional orthogonality conditions:

$$
\begin{aligned}
\left(c_0 + c_{erm} r_{erm} + c_{FM} r^{FM}\right) r_{erm} - r_{erm} + \lambda^{erm} &= 0, \\
\left(c_0 + c_{erm} r_{erm} + c_{FM} r^{FM}\right) r^{FM} - r^{FM} + \lambda^{FM} &= 0,
\end{aligned}
\tag{6}
$$

which we can incorporate into the moment conditions given by (5) to produce efficient estimates of the SDF loadings (c_0, c_{erm}, and c_{FM}), the prices of risk λ^{erm} and λ^{FM}, and their associated standard errors (see Li, Vassalou and Xing (2006)). As testing portfolios we use the 20 state beta-sorted portfolios from Table 1 Panel B of this Appendix plus the twenty-five Fama-French portfolios sorted first by size from smaller to larger and sorted then within each quintile by book-to-market from lower to higher. The system involves $N = 20 + 25 + 2 = 47$ orthogonality conditions to estimate $L = 5$ parameters.

In Table 3 of this Appendix, the KEEPM Model (second row) is supported by the data, with a sizeable quarterly premium (-2.045) significant at almost the 1% level. The corresponding coefficient in the pricing kernel, c_{FM}, is also significant at the 1% level. Looking at the KEEPM model extended with the Fama and French factors (bottom row), all the estimated prices of risk are very similar to those reported in Table 3 Panel A with a quarterly premium on the orthogonal labor risk factor equal to -0.936, significant at the 5% level.

We estimate next the prices of risk from the KEEPM model per division. Table 4 of the Appendix presents the coefficients and prices of risk estimated in each division using the orthogonality conditions in (5) and (6). When comparing these estimates and their statistical significance with those in Table 4 in the paper, obtained using the Fama-MacBeth methodology, we find, with the exception of MA and EN which have a larger estimates, very similar results in terms of the size and the extent of the statistical significance of the estimate prices of risk.

Summarizing, the results regarding the statistical significance of the estimated prices of risk are robust to two different estimation techniques, the traditional Fama-MacBeth cross sectional regressions (reported in tables 2 of the Appendix and 4 of the paper) and the GMM technique (reported in tables 3 and 4 of the Appendix) which is robust to heteroscedacticity.

References

[1] Cochrane, J.H., 2005, Asset Pricing, Princeton University Press.

[2] Li, Q. Vassalou, M., and Y. Xing, 2006, Sector Investment Growth Rates and the Cross-Section of Equity Returns, Journal of Business 79, 1637–1665.

Table 1
Beta-sorted portfolios
Time-series regressions: Aggregate

US stocks are sorted according to their slope coefficients $\hat{\beta}_i^F$ estimated in Table 1 in the paper into three, ten and twenty equally weighted portfolios denoted by subscript p. The quarterly return on these portfolios is estimated over the following year. The time series of quarterly portfolio returns starts in 1965Q1 and ends in 2011Q4.

Panel A reports the average percentage quarterly return on the difference between the portfolio that includes the stocks with the highest betas (P1) and the portfolio with the lowest betas (P3, P10 and P20, respectively). Absolute t-statistics that test whether the difference between the two portfolios is different from zero are reported in parenthesis.

We next estimate an aggregate factor mimicking (FM) portfolio for the orthogonal state labor income risk by going long on the top portfolio containing one-third of the stocks with the highest beta (P1) and short on the bottom portfolio containing one-third of the stocks with the lowest beta (P3). Let r_t^{FM} denote the return on the factor mimicking portfolio. Panel B reports the average returns on the 20 beta-sorted portfolios, the correlation coefficient ρ between each portfolio return and the return on the factor mimicking portfolio, and the coefficient from the following regression (absolute t-values in parenthesis):

$$r_{p,t} = \alpha_p + \beta_p^{FM} r_t^{FM} + \beta_p^{erm} r_{erm,t} + u_{p,t}.$$

Panel A: Average return spread

$P1 - P3$	$P1 - P10$	P1-P20
-1.078 (2.45)	-1.333 (2.11)	-1.294 (1.79)

Panel B: Portfolio statistics

	r_p	$\rho(r_p, r^{FM})$	$\hat{\beta}_p^{FM}$	α_p
P1	2.033 (2.25)	-0.469	-0.142 (1.43)	0.004 (0.79)
P2	2.108 (2.84)	-0.480	-0.064 (0.97)	0.007 (2.02)
P3	2.006 (2.74)	-0.498	-0.113 (1.69)	0.006 (1.70)
P4	2.101 (2.97)	-0.519	-0.136 (2.28)	0.007 (2.23)
P5	1.840 (2.60)	-0.561	-0.234 (3.90)	0.004 (1.23)
P6	1.867 (2.64)	-0.584	-0.282 (4.81)	0.004 (1.22)
P7	2.185 (2.85)	-0.607	-0.378 (5.76)	0.005 (1.63)
P8	2.444 (3.27)	-0.610	-0.374 (5.89)	0.008 (2.56)
P9	2.257 (3.32)	-0.653	-0.497 (7.81)	0.009 (2.60)
P10	2.511 (3.32)	-0.672	-0.517 (9.31)	0.007 (2.59)
P11	2.307 (2.85)	-0.687	-0.600 (9.64)	0.005 (1.45)
P12	2.545 (3.13)	-0.705	-0.659 (10.56)	0.006 (2.08)
P13	2.446 (2.66)	-0.742	-0.855 (13.21)	0.002 (0.83)
P14	2.875 (3.25)	-0.720	-0.766 (11.48)	0.008 (2.34)
P15	3.070 (3.43)	-0.760	-0.919 (14.32)	0.008 (2.63)
P16	2.984 (2.97)	-0.789	-1.111 (16.85)	0.005 (1.58)
P17	3.230 (3.17)	-0.791	-1.122 (17.36)	0.007 (2.21)
P18	2.705 (2.53)	-0.790	-1.191 (16.74)	0.001 (0.31)
P19	3.428 (3.04)	-0.820	-1.387 (19.64)	0.006 (1.74)
P20	3.327 (2.51)	-0.810	-1.653 (16.65)	0.001 (0.23)

Table 2
Beta-sorted and Fama and French portfolios
Cross-sectional Fama-MacBeth regressions: Aggregate

Panel A reports estimates from Fama-MacBeth cross-sectional regressions at each period t:

$$r_p = \lambda_t^0 + \sum_f \lambda_t^f \hat{\beta}_p^f + \lambda_t^{erm} \hat{\beta}_p^{erm} + \xi_p,$$

across portfolios p. As testing portfolios we use the 20 state beta-sorted portfolios from Table 1 Panel B plus the twenty-five Fama-French portfolios sorted first by size from smaller (FF1) to larger (FF21) and sorted then within each quintile by book-to-market from lower to higher. $\hat{\beta}^{erm}$ is the estimated beta for the market risk factor; $\hat{\beta}^f$ is the estimated beta for the risk factor f. In each specification, we estimate and report the average cross-sectional (percentage) quarterly prices of risks. In the first row we estimate the market portfolio price of risk, λ^{erm}, from the CAPM. In the second row we add the price of risk of the state labor income factor mimicking portfolio, λ^{FM}, from the KEEPM. In row number three we estimate the Fama and French (FF) three-factor model including the the size (λ^{smb}) and book-to market (λ^{hml}) prices of risk. In the final row we test the KEEPM augmented with the Fama-French factors (KEEPM-FF). $R^2 = \left[Var_c(\overline{r}_p) - Var_c(\overline{e}_p)\right]/Var_c(\overline{r}_p)$, where Var_c is the cross-sectional variance, $\overline{r}_p$ is the average return and $\overline{e}_p$ is the average residual. $\overline{R}^2$ is the adjusted R^2. Absolute t-values are reported in parenthesis

Panel B presents the square root of the squared pricing errors for the Fama and French (Panel B.1) and state beta-sorted (Panel B.2) testing portfolios and the four model specifications. * indicates that the individual pricing error is statistically significant at the 5% level. Panel B.3 includes the average cross-sectional pricing errors from each model for the forty-five portfolios from panels B.1 and B.2. We also report a test of whether the pricing errors are jointly zero. This is a Chi-sq test given as $\widehat{\alpha}' cov(\widehat{\alpha})^{-1}\widehat{\alpha}$, where $\widehat{\alpha}$ is the vector of average pricing errors across the forty-five portfolios and cov is the covariance matrix of the pricing errors. p-values in brackets.

Panel A: Prices of risk

λ^0	λ^{FM}	λ^{erm}	λ^{smb}	λ^{hml}	$\overline{R}^2$
4.773	−0.880	−3.054			0.66
(3.93)	(2.23)	(2.32)			
1.273		0.935			0.08
(1.46)		(0.85)			
1.023		0.323	0.903	1.283	0.73
(1.12)		(0.29)	(2.04)	(2.83)	
2.128	−0.896	−0.700	0.745	1.224	0.75
(2.37)	(1.98)	(0.64)	(1.74)	(2.75)	

Panel B: Pricing errors

	CAPM	KEEPM	FF	KEEPM-FF
Panel B.1: Fama-French portfolios				
FF1	1.871*	1.220*	1.290*	1.206*
FF2	0.016	0.035	0.021	0.003
FF3	0.185	0.328	0.142	0.197
FF4	0.857*	0.131	0.304*	0.246
FF5	1.115*	0.501*	0.167	0.221
FF6	1.141*	0.003	0.233	0.076
FF7	0.200	0.028	0.086	0.011
FF8	0.479*	0.568*	0.311	0.413*
FF9	0.609*	0.394	0.161	0.226
FF10	0.726*	0.484	0.132	0.022
FF11	1.057*	0.090	0.110	0.186
FF12	0.076	0.348	0.258	0.316
FF13	0.109	0.071	0.028	0.053
FF14	0.401	0.401	0.087	0.169
FF15	0.981*	0.510	0.289	0.319
FF16	0.680	0.419*	0.692*	0.684*
FF17	0.670*	0.080	0.249	0.168
FF18	0.161	0.147	0.015	0.035
FF19	0.218	0.386	0.187	0.239
FF20	0.252	0.354	0.170	0.050
FF21	0.995*	0.313	0.528*	0.310
FF22	0.724*	0.210	0.115	0.035
FF23	0.774*	0.531*	0.216	0.300
FF24	0.547*	0.544*	0.393*	0.461*
FF25	0.504	0.524	0.506*	0.549

Panel B.2: Labor income sorted portfolios

P1	0.312	0.4801	0.215	0.061
P2	0.121*	0.3468	0.076	0.203
P3	0.201	0.0698	0.160	0.054
P4	0.092	0.0680	0.018	0.067
P5	0.347*	0.4152*	0.422*	0.397*
P6	0.316*	0.4992*	0.370*	0.406*
P7	0.061	0.1835	0.178	0.175
P8	0.219	0.0152	0.111	0.079
P9	0.322	0.0213	0.080	0.036
P10	0.231	0.0423	0.166	0.087
P11	0.008	0.3622	0.139	0.223
P12	0.252	0.2663	0.050	0.077
P13	0.024	0.3647	0.255	0.334
P14	0.494	0.1178	0.168	0.119
P15	0.663*	0.0759	0.322	0.204
P16	0.480	0.1028	0.115	0.016
P17	0.701*	0.1928	0.366	0.258
P18	0.128	0.3224	0.217	0.305
P19	0.803*	0.1452	0.548	0.372
P20	0.536	0.0165	0.148	0.034

Panel B.3: Average pricing errors

	CAPM	KEEPM	FF	KEEPM-FF
Average p.e.	0.483	0.287	0.241	0.222
p.e. Test	61.132 [0.04]	74.442 [0.00]	55.182 [0.08]	72.521 [0.00]

Table 3
GMM Cross-Sectional Tests of KEEPM: Aggregate tests

This table presents the GMM estimates of the following pricing equation:

$$E[r_p \left(c_0 + c_{erm} r_{erm} + c_{_{FM}} r^{FM}\right)] = 0,$$

where r_p is the excess return on portfolio p; c_0 is the intercept; c_{erm} is the loading on the market factor; r_{erm} is the return on the market portfolio; $c_{_{FM}}$ is the loading corresponding to the aggregate (common across divisions) orthogonal component of local labor income; and r^{FM} is the return on the orthogonal labor factor mimicking portfolio. In the table, λ^{FM} is the quarterly risk premium corresponding to the local orthogonal labor income factor, and λ^{erm} is the quarterly risk premium corresponding to the market factor. c_{smb} and c_{hml} denote, respectively, the SMB and HML Fama-French factor loadings in the pricing kernel. λ^{smb} and λ^{hml} are the corresponding prices of risk. Absolute t-values in parenthesis. p-values between brackets.

c_0	$c_{_{FM}}$	c_{erm}	c_{smb}	c_{hml}	λ^{FM}	λ^{erm}	λ^{smb}	λ^{hml}
0.956	5.746	4.946			−2.045	−0.858		
(33.06)	(5.68)	(4.63)			(2.68)	(2.20)		
0.943		1.669				−1.095		
(76.48)		(1.75)				(1.41)		
1.056		1.718	−5.181	−4.569		−0.581	1.011	1.434
(24.86)		(1.46)	(3.72)	(3.74)		(0.78)	(2.79)	(3.42)
1.014	1.973	3.283	−4.035	−3.251	−0.936	−1.190	0.936	1.309
(24.97)	(1.29)	(2.57)	(2.29)	(2.51)	(2.27)	(1.53)	(2.30)	(3.06)

Table 4
GMM Cross-Sectional Tests of KEEPM: Per Division

This table presents the GMM estimates of the following pricing equation:

$$E[r_{p,k} \left(c_{0,k} + c_{erm,k} r_{erm,k} + c_k r_k^{FM}\right)] = 0,$$

where $r_{p,k}$ is the excess return on portfolio p in division k; $c_{0,k}$ is the intercept; $c_{erm,k}$ is the loading on the market factor; $r_{erm,k}$ is the return on the market portfolio; c_k is the loading corresponding to the orthogonal component of local labor income in division k; and r_k^{FM} is the return on the orthogonal labor factor mimicking portfolio in division k. In the table, λ_k^{FM} is the quarterly risk premium corresponding to the local orthogonal labor income factor, and λ_k^{erm} is the quarterly risk premium corresponding to the market factor, all per division k. Absolute t-values in parenthesis. p-values between brackets.

	$c_{0,k}$	c_k	$c_{erm,k}$	λ_k^{FM}	λ_k^{erm}
MA	1.084	−2.248	−6.889	−2.045	4.688
	(17.41)	(1.17)	(2.18)	(2.68)	(2.45)
NE	0.976	0.950	−0.283	−0.491	0.479
	(24.10)	(0.55)	(0.12)	(1.00)	(0.31)
SA	0.990	1.500	0.025	−1.167	0.564
	(23.72)	(0.93)	(0.01)	(1.82)	(0.32)
EN	1.037	−0.379	−3.350	−0.491	2.616
	(20.48)	(0.25)	(1.43)	(1.16)	(1.57)
PA	1.027	−1.646	1.194	−1.703	1.734
	(21.97)	(0.55)	(0.86)	(2.48)	(0.91)
ES	0.997	−0.826	1.014	−1.300	1.033
	(29.37)	(0.75)	(0.35)	(1.48)	(0.48)
WS	1.010	1.789	−0.607	−1.673	0.910
	(23.41)	(1.86)	(0.25)	(2.48)	(0.54)
WN	1.025	−0.515	−3.029	−0.517	2.191
	(17.73)	(0.26)	(0.93)	(1.00)	(1.08)
MO	1.201	1.520	−0.736	−1.697	0.819
	(17.35)	(1.87)	(0.24)	(2.24)	(0.37)

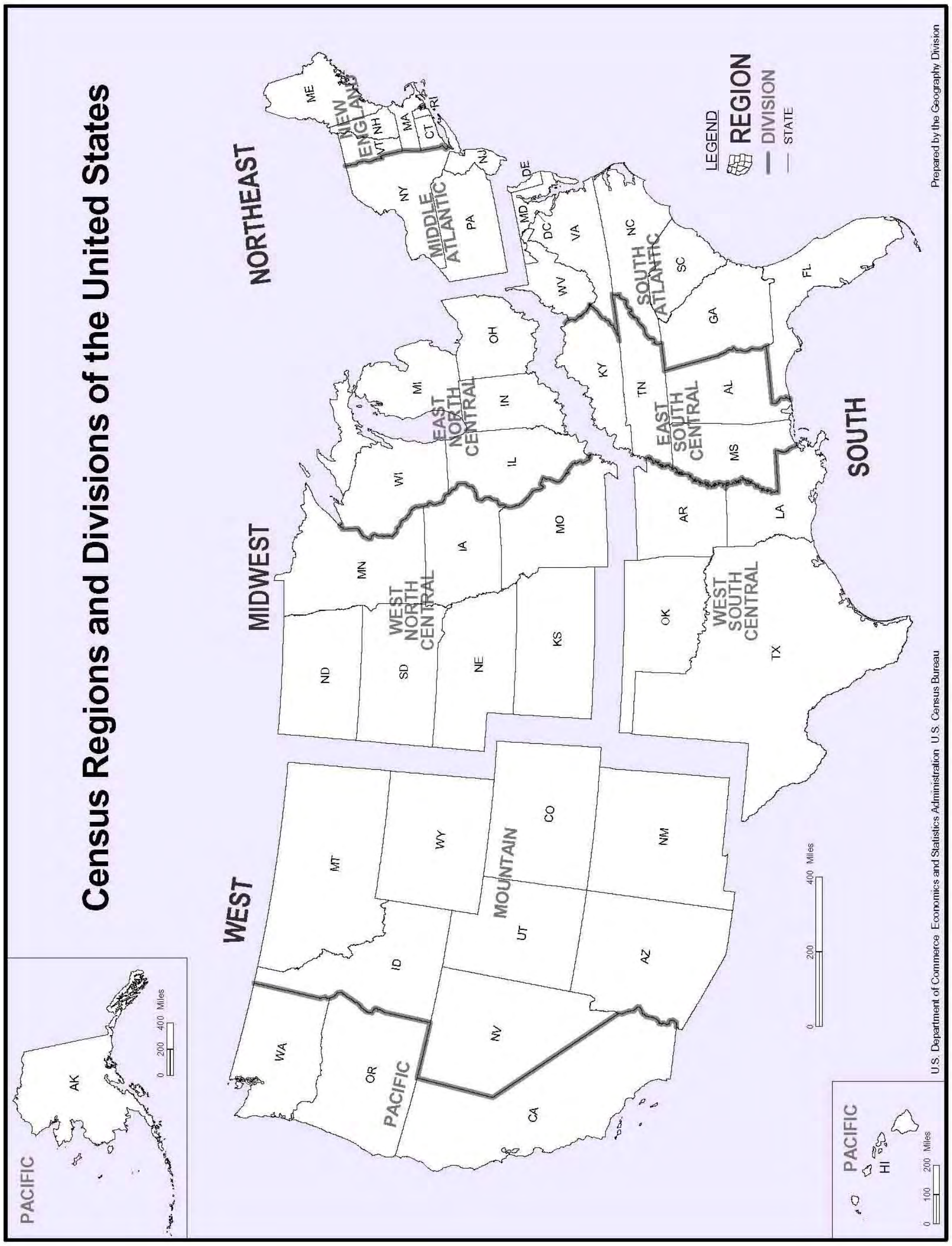

Figure 1: US Census Regions and Divisions. We use the nine census divisions, that we denote with two letters: West South Central (WS), Pacific (PA), East South Central (ES), Mountain (MO), East North Central (EN), South Atlantic (SA), West North Central (WN), Middle Atlantic (MA), and New England (NE).

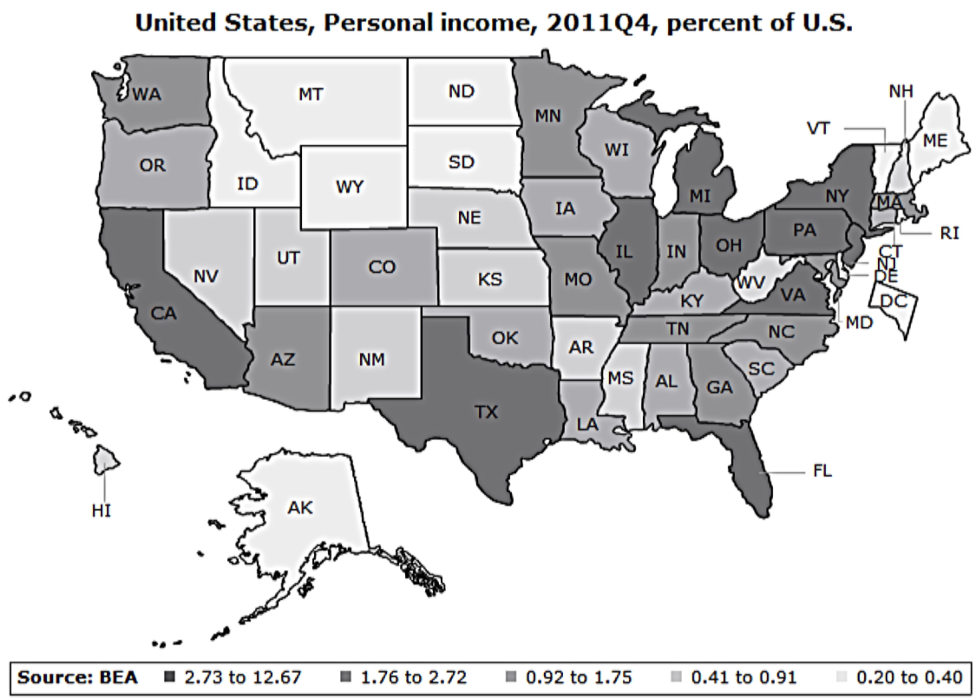

Figure 2: Personal Income per state as a percentage of US total personal income. We use this as a proxy of non-diversifiable income to study if, as predicted by the model, the size of the risk premium is related to the amount of non-diversifiable labor income.