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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 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 and 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 US 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 of the divisions (some states have low 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 three-factor Fama and French (1992) model. In terms of pricing errors and R-square, 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

[1] The US Census partitions the country in nine divisions (see Figure 1 in the Appendix).

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 (2014) 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 US (the four US census regions and eight 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 (2008) and Bernile, Korniotis, Kumar and Wang (2011) 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 US 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 US 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 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 de-

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

nominations represent country divisions indexed by k (we will use both states and US 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:

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

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:

$$r_k^w = \beta_k \, r_M + r_k^F. \tag{2}$$

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 (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

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

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 (3) can be re-written as $E(r) = \Omega\, \boldsymbol{F W}$. 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:

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

$$(5) \qquad \lambda^k = -H\left(\omega_k\theta_k\, b_k\, \mathrm{Var}(r_k^F)+\sum_{k'\neq k}\omega_{k'}\theta_{k'}\, b_{k'}\, \mathrm{Cov}(r_k^F, r_{k'}^F)\right).$$

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 (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 $\mathrm{cov}(r_k^F, r_{k'}^F) > 0$ 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 ($\mathrm{Cov}(r_k^F, r_{k'}^F) = 0$) 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)

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

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, $N \times (K+1)$, 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 (4), (4) 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 Bureau of Economic Analysis (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 (12) 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 US 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 US 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 three-factor Fama and French 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 US 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

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

and re-estimating the orthogonal beta and the market beta until we have thirty-six 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 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

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

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

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 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 US 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.

[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 upon request. We thank the referee for suggesting this test.

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 US 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, $\hat{\beta}_i^{\text{ERM}}$, 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:

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

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 (7) for every stock in the US, 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 $\hat{\beta}_i^{\text{ERM}}$, respectively, in (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 US 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 $r_{k,t}^{\text{FM}}$ 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 one 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 zero 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, 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

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

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, the spreads increase in (absolute) size in all divisions.

We report the excess returns for each portfolio in Panel C. 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 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:

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

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 Table 4 Panel A 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 NE 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 Table 3 Panel C. In all cases the test rejects the null hypothesis that 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 US 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

[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.

[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.

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 (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 Section C. According to the BEA (a map is included in the Appendix, figure 2) there is a high concentration of non-diversifiable wealth (proxied in our tests by personal income) 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, Table 4 Panel B 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. 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

[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 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 an anonymous referee for suggesting this robustness test.

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

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

The prices of risk of the orthogonal factor mimicking portfolios λ_k^{FM} are very similar to the estimates from the KEEPM reported in Panel A, 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 two Fama and French factors are generally not statistically significant.

The average pricing errors are similar in size to those reported in Panel A 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 US 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 (4) and (5) link explicitly the prices of risk to the deep parameters of the model. We derive the following orthogonality conditions from (4)and (5):

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

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

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 $b = \frac{\gamma}{1-\gamma}$.[13] r_M and r_k^{FM} denote, respectively, the time series of the US 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 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

[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.

outlined in Hansen (1982) provides a natural way to estimate the deep parameters of our model. The system of equations in (12) 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 (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.

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 (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 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

[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.

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 Table 4 Panel A must come, according to (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 US country labor income return. Gómez, Priestley and Zapatero (2009) show that the orthogonal US 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 US we estimate the rolling betas with respect to the orthogonal country labor income return and the US stock market return:

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

The slope coefficients $\hat{\beta}_i^C$ and $\hat{\beta}_i^{\text{ERM}}$ are estimated for every stock in the US. 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:

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

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 US 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 (thirty-six quarters):

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

We then estimate the cross sectional regression every quarter:

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

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 $r^F_{s,t}$ 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 $\hat{\beta}^F_i$ is estimated for every stock in the US. 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 US individual stock returns on their estimated betas:

$$r_i = \lambda^0_t + \lambda^F_t \hat{\beta}^F_i + \lambda^{\mathrm{RM}}_t \hat{\beta}^{\mathrm{RM}}_i + \xi_i.$$

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

Average prices of risk

λ^0	λ^F	λ^{RM}
1.933 (4.64)	−0.198 (2.27)	1.403 (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), NE is 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:

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

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 US, regardless of the location of their headquarters.

Average prices of risk

	Panel A			Panel B		
	Only stocks within the division			All stocks 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 US, 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 $r^{\text{FM}}_{k,t}$ 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):

$$r_{p,k,t} = \alpha_{p,k} + \beta^{\text{FM}}_{p,k} r^{\text{FM}}_{k,t} + \beta^{\text{ERM}}_{k,p} r_{\text{ERM,t}} + u_{p,k,t}.$$

Average return spread

	Panel A Only stocks within the division State labor income		**Panel B** All stocks across divisions 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: $\hat{\beta}^{\text{FM}}_{p,k}$

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

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

across portfolios p in each division k. As testing portfolios we use the ten state beta-sorted portfolios from Table 3 Panel C. $\hat{\beta}^{\mathrm{FM}}$ is the estimated beta for the factor mimicking portfolio from Table 3 Panel D; $\hat{\beta}^{\mathrm{ERM}}$ 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) amongst 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^0_k	λ^{FM}_k	λ^{ERM}_k	$\overline{R}^2$	*p.e.*	*p.e. Test*
MA	2.406 (1.51)	−0.903 (2.01)	−0.412 (0.24)	0.57	0.261	13.120 [0.11]
NE	2.055 (1.03)	−0.500 (1.11)	0.463 (0.24)	0.40	0.190	3.664 [0.89]
SA	2.059 (1.03)	−0.991 (1.73)	0.025 (0.01)	0.77	0.196	3.427 [0.90]
EN	−0.216 (0.13)	−0.101 (0.28)	2.291 (1.28)	0.17	0.253	11.117 [0.20]
PA	0.449 (0.21)	−1.812 (2.63)	1.545 (0.75)	0.76	0.366	0.079 [1.00]
ES	1.117 (0.88)	−1.456 (0.60)	1.046 (0.64)	0.89	0.103	7.266 [0.51]
WS	2.795 (1.31)	−1.805 (2.68)	−0.140 (0.07)	0.92	0.174	2.219 [0.97]
WN	2.366 (1.26)	−0.413 (0.89)	−0.316 (0.14)	0.34	0.132	2.766 [0.95]
MO	−0.546 (0.19)	−1.193 (1.77)	2.199 (0.79)	0.63	0.302	2.979 [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 (4):

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

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 US 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 (0.0103)	0.996 (0.000488)	−0.0421 (0.00940)	0.989 (0.00144)	−0.0194 (0.00864)	0.978 (0.00282)	183.9 [0.000]

Table 6

Individual stocks

Country-wide orthogonal income

Let r_t^C denote the orthogonal country labor income return in the US in period t; $r_{\text{ERM},t}$ 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:

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

The slope coefficients $\hat{\beta}_i^C$ and $\hat{\beta}_i^{\text{ERM}}$ are estimated for every stock in the US. 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:

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

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 (thirty-six quarters):

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

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 US 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)