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Human capital and international portfolio diversification: A reappraisal[☆]

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ABSTRACT

We study the implications of human capital hedging for international portfolio choice. First, we document that, at the household level, the degree of home country bias in equity holdings is increasing in the labor income to financial wealth ratio. Second, we show that a heterogeneous agent model in which households face short selling constraints and labor income risk, calibrated to match both micro and macro labor income and asset returns data, can both rationalize this finding and generate a large aggregate home country bias in portfolio holdings. Third, we find that the empirical evidence supporting the belief that the human capital hedging motive should skew domestic portfolios toward foreign assets, is driven by an econometric misspecification rejected by the data.

"Human wealth is likely to be about two-thirds of total wealth and twice financial wealth. This suggests that the omission of human wealth may be a serious matter."

[Campbell (1996)]

1. Introduction

International finance theory emphasizes the effectiveness of global portfolio diversification strategies for cash-flow stabilization and consumption risk sharing.[2] However, the empirical evidence on international portfolio holdings favors a widespread lack of diversification across countries and a systematic bias toward home country assets (see, e.g., Coeurdacier and Rey, 2013 for a recent survey). This discrepancy between theoretical predictions and observed portfolio constitutes the international diversification puzzle (see, e.g., French and Poterba, 1991).

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2 Nevertheless, the size of gains from international risk sharing continues to be a debated issue. E.g.: Grauer and Hakansson (1987) suggest that an individual's gains from international stock-portfolio diversification are large; Cole and Obstfeld (1991) find small gains from perfect pooling of output risks; Obstfeld (1994) calibration exercises imply that most countries reap large steady-state welfare gains from global financial integration; Palacios-Huerta (2001) finds that, for a mean variance investor, adding human capital to the definition of wealth generates substantially smaller gains from international portfolio diversification.

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Moreover, albeit the degree of home bias has been reducing during the last few decades, it remains a first order characteristic of portfolio holdings.[3]

In the major industrialized countries, roughly two-thirds of gross domestic product goes to labor and only one-third to capital. Thus, human wealth likely constitutes about two-thirds of total wealth, suggesting that if investors attempt to hedge against adverse fluctuations in returns to human capital when making financial investment decisions, the mere size of human capital in total wealth makes its potential impact on portfolio holdings self-evident.

Based on this observation, several contributions have argued that when the role of human capital is explicitly taken into account, the observed home country bias in portfolio holdings becomes harder to rationalize. The argument, originally formalized by Brainard and Tobin (1992) with a stylized example, works as follows: if returns to human capital are more correlated with the domestic stock market than with the foreign ones, labor income risk can be more effectively hedged with foreign assets than with domestic ones, and equilibrium portfolio holdings should be skewed toward foreign securities.[4] As emphasized by Cole (1988), “this result is disturbing, given the apparent lack of international diversification that we observe.”

However, as first suggested in Bottazzi et al. (1996), human capital hedging could also lead toward home country bias in portfolio holdings. For instance, the correlation between domestic return on physical and human capital can be lowered by idiosyncratic shocks that lead to a redistribution of total income between capital and labor. In the presence of these rent shifting shocks, foreign assets become a less attractive hedge for labor income risk—especially if total factor productivity shocks are highly correlated internationally. If the size of the rent shifting shocks is large enough, a situation in which domestic assets are the best hedge against human capital risk arises, therefore leading to home country bias in portfolio holdings.

In this paper, we ask whether the human capital hedging motive is likely to have a sizeable effect on optimal portfolio choice, and what its implications are for the international diversification puzzle. Moreover, we propose a rationalization of the home country bias, based on a setting of endogenous portfolio formation and incomplete markets, that not only can rationalize aggregate portfolio holdings, but also the variable degree of home country bias in households' portfolios.

In particular, Fig. 1 depicts a novel (to the best of our knowledge) finding about households' equity home bias.[5] Panel (a) depicts the (locally weighted regression of the) share of foreign assets in U.S. household portfolios as a function of the household financial wealth to labor income ratio. Since (labor income) flows and stocks (of human capital) are cointegrated, Panel (a) shows that there is a systematic relationship between household specific home country bias and the household specific financial to human capital ratio: the degree of home country bias monotonically decreases as the human capital component of household total wealth becomes smaller relative to the household financial wealth. That is, in micro data, when the human capital hedging motive is more prominent relative to the financial wealth hedging motive, household portfolios show a higher degree of home country bias. Moreover, panel (b) of Fig. 1, that depicts the (locally weighted regression of the) number of stocks in U.S. household portfolios as a function of the household financial wealth to labor income ratio, shows that when the households' human capital wealth is relative larger than the financial wealth, the household portfolio will tend to be overall less diversified.

Our paper provides a rationalization of both of these findings, as well as of the aggregate home country bias, and also shows that the canonical intuition that human capital should skew portfolio holdings toward foreign assets, and the related supporting empirical evidence, are both very fragile. In particular, we offer two main contributions.

First, using novel estimates of the correlations of human capital and stock market return innovations, we calibrate an incomplete market model in which agents face both idiosyncratic and aggregate labor income risk, as well as borrowing constraints.[6] The model is also calibrated to match the microeconomic (following Gourinchas and Parker, 2002) characteristics of the U.S. labor income and track the distribution of the asset wealth to labor income ratios observed in the Panel Study of Income Dynamics (PSID).

The main findings of this calibration exercise are that *a*) investors that enter the stock market with a low level of liquid (i.e. financial) wealth to labor income ratio will initially specialize in domestic assets and, *b*) only as the level of asset wealth to labor income ratio increases do agents start diversifying their portfolios internationally by progressively adding different assets to their holdings, *c*) as a consequence, the aggregate portfolio of U.S. investors shows a large degree of home bias.

What drives these results? Households face large human capital risk, but this is mostly of the idiosyncratic type—hence underestimated in a homogeneous agent setting. Moreover, in the presence of liquidity constraints, agents cannot borrow to construct an optimally diversified portfolio. Therefore, when their level of liquid wealth to labor income ratio is sufficiently high and they enter the stock market, agents try to minimize the overall wealth risk, investing first in the asset that has the lowest degree of correlation with labor income innovations—and, as discussed below, this assets is, in the data, the domestic stock. Only when the ratio of liquid wealth to labor income is sufficiently high, and the labor income risk hedging motive becomes less important relative to the financial risk hedging one, do agents start investing in foreign assets and diversifying their portfolios internationally. Since the distribution of liquid wealth to labor income is (in the data as in the model) concentrated in the region of low liquid wealth to labor income ratios, the resulting aggregate portfolio is heavily skewed toward domestic assets. Note that, in the absence of market frictions and idiosyncratic risk, the estimated and calibrated correlations of labor income and returns innovations, being very small, would have almost no effect on the optimal portfolios.

Moreover, since in our model the aggregate home country bias depends on both the household optimal investment policy functions and the aggregate distribution of liquid wealth to labor income, a trend of increasing concentration of financial wealth (as documented by Piketty, 2014), and/or a negative trend in the labor share of income (as documented by Karabarbounis and Neiman, 2013), would both generate a negative trend in the degree of home country bias as found by Coeurdacier and Rey (2013).[7]

[3] Coeurdacier and Rey (2013, Table 1) estimate that, for a large set of countries, the 2008 portfolio share in domestic equity is on average about 71%, with an average implied home bias (measured as one minus the ratio of domestic equities in the domestic portfolio relative to the the domestic share in world market capitalization) of about 70%.

[4] Moreover, Michaelides (2003) shows with a calibration exercise that, in the presence of liquidity constraints, if labor income shocks are positively correlated with the domestic stock market returns and orthogonal to foreign asset returns, investors should hold only foreign assets in their portfolios.

[5] We are thankful to Laura Bottazzi and Şebnem Kalemli-Özcan for suggesting us to explore this dimension of the data.

[6] We focus on household level liquidity constraint given the widespread empirical support for this modeling assumption (see, e.g., Zeldes, 1989; Iacoviello, 2005; Attanasio et al., 2008).

[7] This also implies that the role of globalization and, more broadly, multinationals, for the home country bias, cannot be evaluated without controlling for their impact on the distribution of wealth and income. For instance, if globalization where to reduce the international diversification opportunities, while at the same time increasing the concentration of wealth and/or reducing the labor share of income, its effect on the home country bias would be ambiguous.

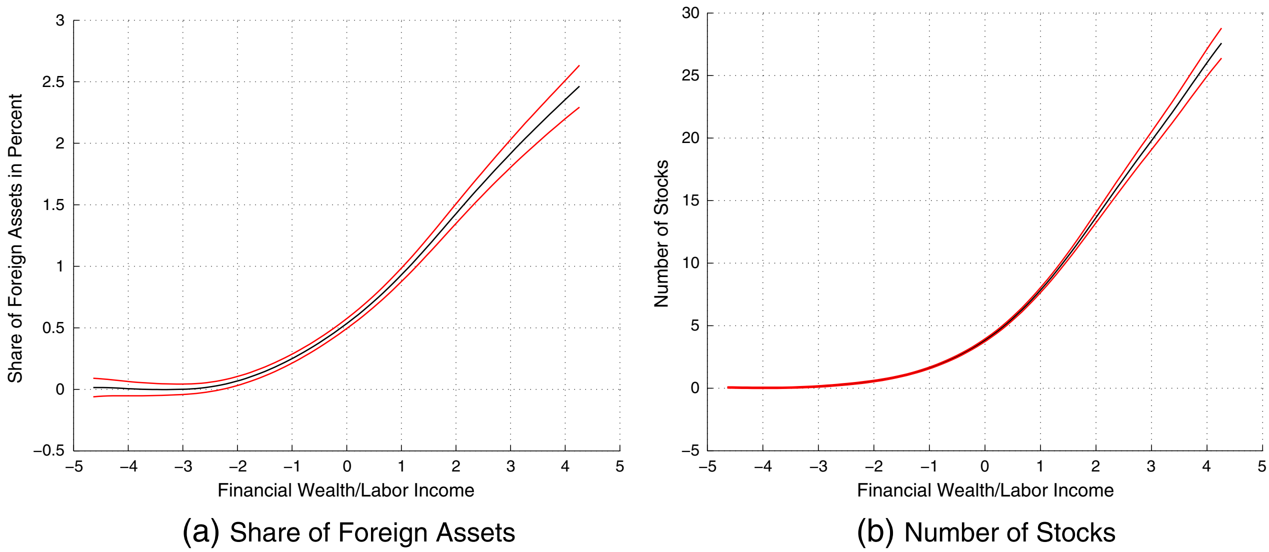

Fig. 1. Locally weighted regression of the share of foreign assets, panel (a), and the number of stocks hold by the household, panel (b), on the logarithm of the financial wealth to labor income ratio of the household. Two standard error confidence bands computed via bootstrapping. Estimation based on all the U.S. Survey of Consumer Finances available to date (1992–2013). Year specific estimates are reported in Fig. A1 of the Appendix.

Since the pattern of correlation of innovations to labor income and returns plays an important role in our calibration, our second contribution hinges upon the identification of a common misspecification that has affected the previous empirical literature, and the provision of novel estimates that are not affected by this issue. In particular, we show that the seminal empirical result of Baxter and Jermann (1997) that, in the presence of a human capital hedging motive, investors should short sell the domestic capital stock—implying that "the international diversification puzzle is worse than you think"—is largely due to an econometric misspecification rejected by the data: the assumption that there are neither cross-country shocks to human and physical capital payoffs, nor common long run trends. We show that, once this restriction is relaxed, the effective degree of technological and economic integration becomes evident, therefore reducing the opportunities to hedge human capital risk by investing in foreign assets.

Moreover, we also show that there is substantial uncertainty attached to the estimation of aggregate physical capital returns via the canonical Campbell and Shiller (1988) cum vector autoregression (VAR) approach. This feature of the data provides a rationalization for the apparently contradictory empirical evidence on the correlation between returns to human and physical capital found in the previous literature (that typically has not reported confidence bands for the estimated correlations): Lustig and Nieuwerburgh (2008) find a strong negative correlation between domestic returns to physical and human capital in U.S. data; Bottazzi et al. (1996) find such a correlation to be negative in all the countries they consider but the United States, where they find it to be strongly positive; Baxter and Jermann (1997) find this correlation to be positive and very close to one for all the countries they consider (United States, United Kingdom, Germany, and Japan).

Nevertheless, we find that by restricting the set of assets available to hedge human capital to include only publicly traded stocks—what we consider as the relevant case for most households—much sharper estimates can be obtained, and these are the estimates used in our calibration exercise discussed above. We find that in this case human capital hedging can help explain the home country bias in portfolio holdings—since domestic returns to human capital tend to be systematically more correlated with foreign stock markets—but, we show, given the small magnitudes of the estimated correlation, the effect is quantitatively very small in a frictionless complete market setting.[8] Nevertheless, as discussed above, this same small correlations have very large aggregate effects in an incomplete market settings in which agents face both idiosyncratic and aggregate labor income risk.

Note that, overall, our result that domestic capital markets constitute a good hedge for domestic human capital risk are in line with the findings of a large empirical literature. Palacios-Huerta (2001) finds that if human capital is included in the definition of wealth, gains from international financial diversification for a mean-variance investor appear to be smaller than previously reported. Lustig and Nieuwerburgh (2008) find that innovations in current and future human wealth returns are negatively correlated with innovations in current and future domestic financial asset returns. Abowd (1989) finds a large and negative correlation between unexpected union wage changes and unexpected changes in the stock value of the firm. Davis and Willen (2000), using data from the PSID to construct synthetic cohorts, find that the correlation between domestic labor income shocks and returns on the S&P 500 is substantially negative for some categories.[9] Moreover, they find that for six out of the eight sex-education groups considered in their study, a long position on the worker's own industry represents a good hedge for labor income risk. The empirical works of Gali (1999), Rotemberg (2003), and Francis and Ramey (2004) also document a negative correlation between labor hours and productivity conditioning on productivity shocks. Coeurdacier and Rey (2013), conditioning on exchange rate movements, find that wages and dividends growth rates comove negatively for all the countries they consider (also Coeurdacier et al., 2013 ; Heathcote and Perri, 2013 provide similar empirical evidence).

Theoretically, we show that a situation in which labor income innovations are more correlated with the domestic payout to capital than the foreign ones—as we find in the data—is likely to arise

[8] This is consistent with the findings of Fama and Schwert (1977).

[9] Davis and Willen (2000) generally find that the degree of correlation between earning shocks and equity returns rises with education, with a lower bound correlation of −0.25 for men who did not finish high school. This is in line with empirical studies on the labor demand in modern economies that consistently find that more educated workers are relatively complementary to physical capital and the use of advanced technologies.

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once the degree of international economic integration observed in the data is properly taken into account. In particular, we show that very small redistributive shocks (shocks with a variance that is equal to as little as 6%–11% of the output variance), are enough to make the domestic equity market the best hedge for human capital risk.

The analysis presented in this paper is part of the literature that has attempted to explain home bias as a hedge against non-tradable risks.[10] Moreover, the potential rationalization of the international diversification puzzle we document in this paper should be interpreted as complementary, rather than alternative, to the ones based on transaction and information frictions (e.g. Van Nieuwerburgh and Veldkamp, 2009; Bhamra et al., 2014), nominal stickiness (e.g., Engel and Matsumoto, 2009), non-traded goods (e.g., Heathcote and Perri, 2013) and more broadly the role of real exchange rate fluctuations (e.g., Coeurdacier, 2009; Kollmann, 2006; Baxter et al., 1998).

The remainder of the paper is organized as follows. Section 2 presents a calibrated model of human capital risk hedging in which households face both idiosyncratic and aggregate labor income risk, as well as liquidity constraints. Section 3 presents the empirical approach undertaken to measure factor returns and rationalizes the difference in results between our findings and the previous empirical literature. The final section outlines the conclusions of the paper, while a detailed data description, as well as additional results and robustness checks, are reported in the Appendix.

2. Portfolio choice with heterogeneous human capital risk and liquidity constraints

In this section we rely on numerical methods to compute the equilibrium outcome of a model that directly takes into account that *i*) most of the human capital risk faced by households is idiosyncratic in nature, and *ii*) households' optimal portfolio choice is influenced by liquidity constraints. The simple incomplete markets model presented below is a generalization of Heaton and Lucas (1997) to a multiple asset context and of Michaelides (2003), and builds upon the household income process estimated by Gourinchas and Parker (2002).

2.1. Model setup and calibration

Each household solves the problem

$$\max_{\left\{C_t, B_t, S_t^d, \left\{S_t^j\right\}_{t=1}^N\right\}} E_0 \sum_{t=0}^{\infty} \beta^t \frac{C_t^{1-\gamma}}{1-\gamma}$$

subject to the short selling constraints B_t, S_t^d, $S_t^j \geq 0$ for all t and j, the period budget constraint

$$C_t + B_t + S_t^d + \sum_{j=1}^{J} S_t^j \leq R_t^f B_{t-1} + R_t^d S_{t-1}^d + \sum_{j=1}^{J} R_t^j S_{t-1}^j + Y_t, \quad (1)$$

and the standard transversality condition, where $1 > \beta > 0$ is the time discount factor (calibrated at the value of 0.95 per year), γ is the relative risk aversion coefficient (calibrated at the benchmark value of 3), C_t is consumption, B_t is the dollar amount invested in domestic bonds, S_t^d is the amount invested in the domestic stock, S_t^j is the amount invested in the stock of country j, Y_t denotes the labor income, R^f is the gross risk free rate, R_t^d is the gross return on the domestic stock, and R_t^j is the return on the stock of country j.

[10] See, e.g., Eldor et al. (1988), Stockman and Dellas (1989), Tesar (1993), Baxter et al. (1998), Obstfeld and Rogoff (2001), Serrat (2001), and Pesenti and van Wincoop (2002).

In order to model aggregate labor income (Y^g) dynamics in a parsimonious manner, we searched for a low dimensional ARIMA representation and selected (see Appendix A.2 for details) an MA(2) specification for its log growth rate:

$$g_{t+1} = \log \frac{Y_{t+1}^g}{Y_t^g} = \mu_y + \varepsilon_{t+1} + \theta_1 \varepsilon_t + \theta_2 \varepsilon_{t-1} \quad (2)$$

where $\varepsilon_t \sim \mathcal{N}(0, \sigma_\epsilon^2)$. The individual labor income of agent i is assumed to follow the process

$$Y_t^i = Y_t^g W_t^i U_t^i \quad (3)$$

$$W_t^i = G W_{t-1}^i N_t^i \quad (4)$$

where U_t^i is independent of ε, N, and asset returns, and $\log U_t^i \sim \mathcal{N}\left(-\frac{1}{2}\sigma_u^2, \sigma_u^2\right)$ so that $E\left[U_t^i\right] = 1$, $\log W_t^i$ evolves as a random walk with drift, $\log N_t^i \sim \mathcal{N}\left(-\frac{1}{2}\sigma_n^2, \sigma_n^2\right)$, so that $E\left[N_t^i\right] = 1$, and N is independent of ε and asset returns. This specification corresponds to Gourinchas and Parker (2002) except for the added term Y_t^g that reflects aggregate economic uncertainty.[11] Following Gourinchas and Parker (2002) estimates, we calibrate $\sigma_u = 0.073$ and $\sigma_n = 0.105$, and we calibrate σ_ε, θ_1, θ_2, and μ_y using the point estimates in Table A2 in Appendix A.2. This calibration implies that the aggregate labor risk component has a standard deviation that is of a unit of magnitude smaller than the ones of the idiosyncratic components. We assume also that log returns on risky assets and shocks to the aggregate labor income process (ε) are jointly normal.

Given Eqs. (2)–(4), the individual labor income growth is given by

$$\Delta \log Y_t^i = g_t + \log G + \log N_t^i + \Delta \log U_t^i$$

and requires the normalization $\log G = \frac{1}{2}\left(\sigma_u^2 + \sigma_n^2\right)$ in order to recover the aggregate labor income growth rate as an average of the individual labor income growth rates.

The model implies the following Euler equations

$$C_t^{-\gamma} = \beta R^f E_t\left[C_{t+1}^{-\gamma}\right] + \lambda_B$$

$$C_t^{-\gamma} = \beta E_t\left[C_{t+1}^{-\gamma} R_{t+1}^d\right] + \lambda_d$$

$$C_t^{-\gamma} = \beta E_t\left[C_{t+1}^{-\gamma} R_{t+1}^j\right] + \lambda_j \quad \forall j$$

where λ_B, λ_d, and λ_j are the Lagrange multipliers on the short selling constraints for domestic bonds, domestic stocks, and foreign stocks. Let X_t be the cash-on-hand at the beginning of period t

$$X_t = R^f B_{t-1} + R_t^d S_{t-1}^d + \sum_{j=1}^{J} R_t^j S_{t-1}^j + Y_t.$$

Since the utility function implies that there is no satiation in consumption, the budget constraint will hold with equality and

$$C_t = X_t - 1_{\{B_t > 0\}} B_t - 1_{\left\{S_t^d > 0\right\}} S_t^d - \sum_{j=1}^{J} 1_{\left\{S_t^j > 0\right\}} S_t^i \quad (5)$$

[11] Gourinchas and Parker (2002) also add a small positive probability for $U = 0$, therefore allowing the labor income to be zero with positive probability.

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where $1_{\{.\}}$ is an index function that takes value 1 if the condition in brackets is satisfied and zero otherwise. To solve the model, we make the problem stationary dividing all the variables at time t by

$$Z_t^i := E_t\left[Y_{t+2}^i\right] = G^2 W_t^i Y_t^g \exp\left[(\theta_1+\theta_2)\,\varepsilon_t + \theta_2\varepsilon_{t-1} + k\right]$$

where $k = 2\mu_y + \left[1+(1+\theta_1)^2\right]\frac{\sigma_\varepsilon^2}{2}$. Note also that

$$\log\frac{Z_{t+1}^i}{Z_t^i} = \mu_y + (1+\theta_1+\theta_2)\,\varepsilon_{t+1} + \log G + \log N_{t+1}^i$$

Using Eq. (5) and the homogeneity of degree $-\gamma$ of the marginal utility, we can rewrite the Euler equations as

$$\left(x_t - b_t - s_t^d - \sum_{j=1}^J s_t^j\right)^{-\gamma} = \max\left\{\left(x_t - s_t^d - \sum_{j=1}^J s_t^j\right)^{-\gamma};\ \beta R^f E_t\left[c_{t+1}^{-\gamma}\left(\frac{Z_{t+1}^i}{Z_t^i}\right)^{-\gamma}\right]\right\}$$

$$\left(x_t - b_t - s_t^d - \sum_{j=1}^J s_t^j\right)^{-\gamma} = \max\left\{\left(x_t - b_t - \sum_{j=1}^J s_t^j\right)^{-\gamma};\ \beta E_t\left[R_{t+1}^d c_{t+1}^{-\gamma}\left(\frac{Z_{t+1}^i}{Z_t^i}\right)^{-\gamma}\right]\right\}$$

$$\left(x_t - b_t - s_t^d - \sum_{j=1}^J s_t^j\right)^{-\gamma} = \max\left\{\left(x_t - b_t - s_t^d - \sum_{j=1, j\neq j'}^J s_t^j\right)^{-\gamma};\ \beta E_t\left[R_{t+1}^{j'} c_{t+1}^{-\gamma}\left(\frac{Z_{t+1}^i}{Z_t^i}\right)^{-\gamma}\right]\right\}\quad \forall j' = 1,\ldots J$$

where small font letters represent the ratios of the capitalized variables to the normalizing variable Z (e.g., $c := C/Z$), and the normalized state variable x (see, e.g., Deaton, 1991 ; Carroll, 2009) evolves according to

$$x_t = \left(R^f b_{t-1} + R_t^d s_{t-1}^d + \sum_{j=1}^J R_t^j s_{t-1}^j\right)\frac{Z_{t-1}^i}{Z_t^i} + \frac{Y_t^i}{Z_t^i} \tag{6}$$

where

$$\frac{Y_t^i}{Z_t^i} = G^{-2} U_t^i \exp\left[-(\theta_1+\theta_2)\,\varepsilon_t - \theta_2\varepsilon_{t-1} - k\right].$$

In order for the individual Euler equations to define a contraction mapping for the normalized asset holdings optimal rules $\{b(x,\varepsilon), s^d(x,\varepsilon), s^j(x,\varepsilon)\}$, we need (following Theorem 1 of Deaton and Laroque, 1992) that

$$\beta R_{t+1}^f E_t\left[\left(\frac{Z_{t+1}}{Z_t}\right)^{-\gamma}\right] < 1$$

$$\beta E_t\left[R_{t+1}^d\left(\frac{Z_{t+1}}{Z_t}\right)^{-\gamma}\right] < 1$$

$$\beta E_t\left[R_{t+1}^i\left(\frac{Z_{t+1}}{Z_t}\right)^{-\gamma}\right] < 1.$$

Table 1
Preference and labor income parameters.

γ	3
β	0.95
σ_U	0.210
σ_N	0.146
μ_y	0.019
ϑ_1	0.448
ϑ_2	0.094
Mean market return	0.060
Market return st. dev.	0.175
Risk free rate	0.011

Given the assumptions on the primitives and the calibrated values, these conditions hold and there exists a unique set of optimal policies satisfying the Euler equations.

To avoid the curse of dimensionality of numerical solutions and, most importantly, in order to have sufficiently long time series for the estimation of the variance covariance matrix of labor income innovations and asset returns using the unrestricted VAR approach presented in Section 3 below we focus on four countries: the United States (as domestic country), the United Kingdom, Japan, and Germany. As we show in Section 3 below not restricting the VAR representation to have country specific block exogeneity is both required by the data, and needed in order to uncover the true degree of hedging potential via international diversification.[12]

Since the U.S. domestic risky asset has enjoyed both the lowest variance and the highest Sharpe ratio compared to the other countries considered, and this pushes the optimal portfolio to be skewed toward the domestic stock, we calibrate all the countries as having the same mean return and Sharpe ratio as the United States. A summary of the calibrated preference and labor income process parameters are reported in Table 1.

The crucial element in calibrating the model is the covariance structure of asset returns and innovations to the aggregate labor income process. We measure capital returns using broad stock market indexes and calibrate their covariance using the time series sample analogous. The calibration of the covariance structure of aggregate labor income shocks and stock market returns is summarized by the correlations reported in the first four columns of Table 2. These are based on the estimation approach discussed extensively in Section 3 below where we show that the (different) estimates obtained in the previous literature are due to misspecification. The crucial element in Table 2 is that the correlation between U.S. labor income innovations (fourth column) with the domestic stock market is marginally smaller than the ones with foreign stock markets returns (expressed in dollar terms). Note that these correlations are all small in magnitude and, as shown in Table A5, would have a very small effect on the optimal portfolio choice in a complete markets setting.

As a benchmark, the last column of Table 2 reports the implied optimal portfolio shares of the domestic portfolio absent any human capital hedging motive and shows that, according to the estimated covariance structure of returns, the share of U.S. assets in the U.S. domestic portfolio would be about 25% in the absence of aggregate labor income risk.

2.2. Investors' optimal policy rules and portfolio choice

Having calibrated the model, we can estimate the optimal policy function by standard numerical dynamic programming techniques (see, e.g., Carroll, 1992 ; Haliassos and Michaelides, 2002) to compute the optimal consumption and asset holding rules. Since

[12] Note also that, according to the 2012 World Bank data, these four countries alone account for more than half of the world total market capitalization of listed companies.

Table 2
Market returns and aggregate labor income shock correlations.

	Correlations				Implied market portfolio w.o. labor income risk
	Germany	Japan	U.K.	Aggregate labor income shocks	
U.S.	0.57	0.32	0.72	0.04	25%
Germany		0.46	0.51	0.14	22%
Japan			0.41	0.19	36%
U.K.				0.15	18%

the time t optimal policy rules depend both on the normalized cash-on-hand and on the last labor income shock (ε_{t-1}), we numerically integrate out this last variable to have policy rules as a function of the cash-on-hand only,[13] obtaining the investment rules $\{b(x), s^d(x), s^i(x)\}$. Moreover, from Eq. (5) we can obtain the optimal consumption rule $c(x)$.

Optimal policy rules are plotted, as a function of normalized cash-on-hand, in Fig. 2. Not surprisingly, the optimal consumption policy rule has the same shape as in the buffer stock saving literature, with consumption being equal to cash-on-hand (no saving region) until a target level of cash-on-hand is reached and saving starts taking place. Once the saving region is reached, the consumers specialize in stocks, disregarding bonds. This result, well known in the literature, was originally obtained by Heaton and Lucas (1997) in a domestic portfolio choice settings, and it reflects the implication of the large equity premium for the optimal portfolio choice.

More interestingly, when the consumer enters the saving region, she initially invests only in the domestic stocks and only gradually diversifies her portfolio internationally as the level of cash-on-hand increases. This happens for three reasons. First, only a small buffer stock saving is needed for the agent to protect herself from future labor income shocks. Second, when entering the saving region, the agent prefers to invest in the assets that have the smallest correlation with labor income shocks, in order not to increase her overall level of risk correlated with income. This is due to the fact that, when entering the investment region, almost the entirety of the agent's wealth is in the form of human capital. Hence, for relatively low levels of cash-on-hand, the human capital hedging motive dominates the portfolio diversification motive. As a consequence, the order in which the agents start investing in the different stock markets closely match the inverse rank of the correlations between labor income innovations and asset returns. Third, only for very high levels of liquid wealth to labor income ratio (high x) does the financial portfolio diversification motive become more important than the labor income hedging one, and the agent starts diversifying fully her portfolio. This is due to the fact that, as x increases, so does the non-human capital component of the household wealth, therefore reducing the human capital hedging motive.

Comparing this result with the empirical distribution of cash-on-hand in the PSID data set, less than 1% percent of the population should be investing positive amounts in all four of the assets considered. Moreover, given the positive correlation between normalized cash-on-hand and asset wealth observed in the data, the results imply that only the richest households will be diversifying their portfolio internationally, coherently with the empirical evidence on households' portfolio holdings at the micro level (see, e.g., Jappelli et al., 2001).

[13] Optimal policy functions do not seem to change significantly as a function of past aggregate labor income shocks, mainly due to the very small variance of these shocks compared to the idiosyncratic ones. In particular, policy functions computed assuming a plus or minus two standard deviation shock in aggregate labor income are almost identical to the ones obtained after integrating out this variable.

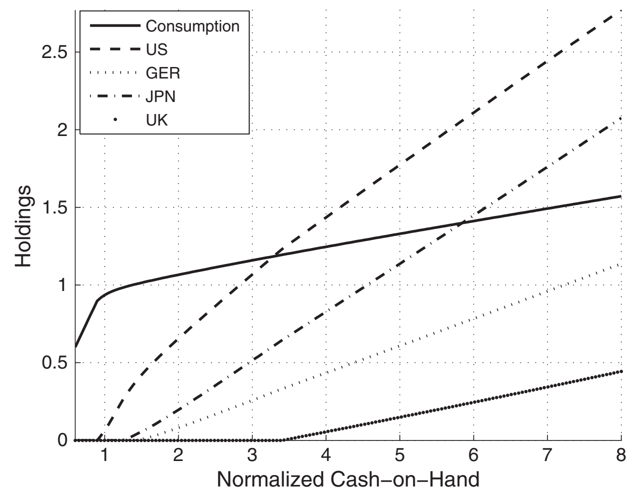

Fig. 2. Optimal consumption and investment policy functions as a function of normalized cash-on-hand.

Using the estimated policy functions, we can compute the optimal portfolio shares as a function of cash-on-hand. These optimal shares are reported in Fig. 3. The figure shows a large bias toward domestic assets in all the relevant ranges of standardized cash on hand, implying that more than 99% of the households should have an asset portfolio strongly biased toward domestic assets. Compared with the optimal share of domestic assets in the market portfolios without aggregate labor income risk (25% in Table 2), this represents a home country bias of individual portfolios that ranges from 75% to 19%. Even investors in the top 1% of the distribution of cash-on-hand observed in the data would have, on average, more than 50% of their asset wealth invested in domestic stocks. Interestingly, this large home bias is generated by extremely small differences in the correlations between labor income shocks and market returns across countries, and a very small aggregate labor income risk component. Moreover, as shown by counterfactual calibration results,[14] this effect is mostly driven by the ordering, rather than the magnitudes, of the correlations between labor income innovations and stock market returns. This implies that small shocks that lower the correlation between aggregate labor income innovations and market returns at the country level can generate, in the presence of short selling constraints and buffer-stock saving behavior, a very large degree of domestic bias in portfolio holdings.

2.3. Implications for the aggregate portfolio

This subsection derives the implications of the optimal investment rules, obtained in the previous subsection, for the aggregate portfolio of U.S. investors.

The standardized cash-on-hand in Eq. (6) follows a renewal process and can be shown to have an associated invariant distribution,[15] and this can be used to compute the implied aggregate portfolio of U.S. investors. Moreover, given the estimated policy functions, the aggregate portfolio can also be computed using the observed empirical distribution of cash-on-hand.

The implied model distribution can be computed in two different ways. First, conditioning on a given value for the lagged aggregate labor income shock (ε_{t-1}), we can use the policy functions and

[14] Available upon request.
[15] See, e.g., Deaton and Laroque (1992), Carroll (1997), Szeidl (2013), and Carroll (2004).

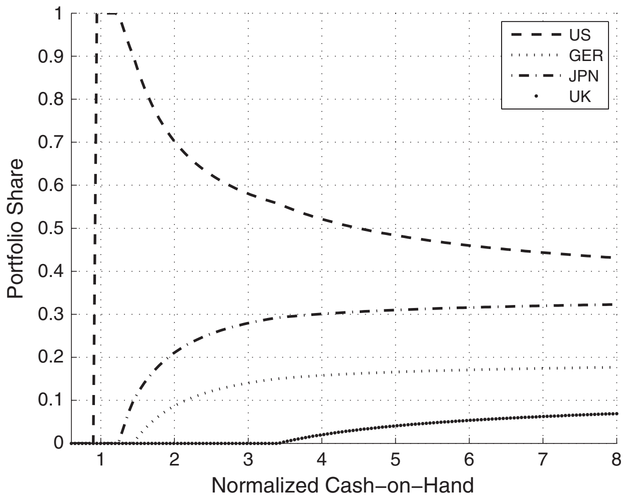

Fig. 3. Optimal portfolio shares as a function of normalized cash-on-hand.

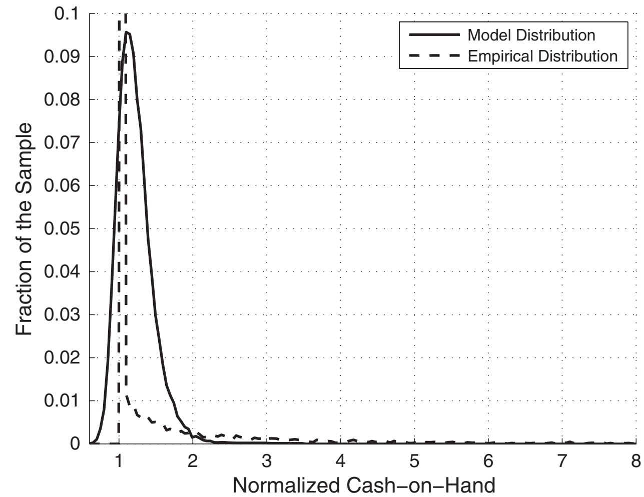

Fig. 4. Model implied and empirical invariant distributions.

Eq. (6) to compute, by repeated simulation over a grid of values, the transition probabilities from one level of cash-on-hand to the other.

$$T_{lm} = \Pr(x = l | x = m).$$

Given the matrix T of transition probabilities, the probability of each state is updated by

$$\pi_{l,t+1} = \sum_m T_{lm} \pi_{m,t}.$$

Therefore, the invariant distribution π can be found as the normalized eigenvector of T corresponding to the unit eigenvalue by solving

$$\begin{pmatrix} T - I & \underline{1} \\ \underline{1}' & 0 \end{pmatrix} \begin{pmatrix} \pi \\ 0 \end{pmatrix} = \begin{pmatrix} 0 \\ 1 \end{pmatrix}$$

where I is the identity matrix and $\underline{1}$ is a vector of ones of appropriate dimension. Since this procedure produces invariant distributions conditioned on the lagged aggregate labor income shock (ε_{t-1}), we can integrate out the conditioning variable to obtain the unconditional *model distribution* of x.

Second, we can alternatively draw random initial levels of x to reproduce the initial heterogeneity in wealth among agents, and then simulate dynamically the evolution of normalized cash-on-hand over time, generating what we refer to as the *dynamic distribution* of the model. We perform both procedures since the first one requires fixing ex-ante the relevant range of x while the second one instead determines the relevant range autonomously, therefore providing a robustness check of the construction of the model ergodic distribution.

Fig. 4 reports the distribution of normalized cash-on-hand implied by the model and the observed distribution of normalized cash-on-hand in the PSID data.[16] The model seems to reproduce fairly well the location of the mode and the shape of the right tail of the empirical distribution, but the model distribution is much less concentrated than the data around the boundary between the saving and the no saving zone, implying a higher participation rate in the market than what is observed in the PSID data, probably due to the absence of stock market entry costs in the setup of the model.

With these distributions at hand, we can compute the implied aggregate portfolio shares of U.S. investors. The first column of Table 3 reports, as a benchmark comparison, the CAPM market portfolio implied by the calibrated covariance structure of returns in the absence of labor income risk. The implied aggregate portfolio shares of the model are reported in the second column. There is a dramatic effect of labor income risk on the aggregate portfolio with about 95% of the market portfolio invested in domestic assets. Moreover, the relative investments in foreign stocks are strongly affected, with a reduction of the portfolio shares in individual foreign stocks moving from the 18%–36% range to the 0%–4% range.

Since agents with different levels of normalized cash-on-hand are likely to have different amounts of wealth invested in the stock market, the simple computation of the aggregate portfolio reported in column two of Table 3 could be a poor approximation of the aggregate portfolio. To address this issue, the third column of Table 3 weights the model distribution by the contribution to the aggregate portfolio of agents having different levels of cash-on-hand. This weighting of the distribution also corrects for the fact that the model implies a higher degree of market participation than what is observed in the data. The weights are constructed from the PSID data and are proportional to the total stock market holdings of households belonging to each category of normalized cash-on-hand.

This weighting somehow reduces the degree of home bias relative to column two, but still delivers a portfolio share of domestic stocks of about 75%, implying that hedging human capital increases the portfolio share of domestic stocks by as much as 50% and decreases the portfolio shares of German, Japanese, and U.K. stocks by, respectively, 15%, 19% and 17%.

The last two columns of Table 3 show that our main result also holds if we compute the aggregate portfolios using the empirical (again, for PSID data), rather than the model implied, distribution of cash-on-hand, with (column five) and without (column four) weighting. The aggregate portfolio shares implied by the empirical distribution are, in both cases, quite similar to the ones obtained by weighting the model distribution (column three) and carry the same message: the human capital hedging motive generates a very large home country bias, with an increase of the portfolio shares of domestic assets between 36% and 50%.

But what is the key mechanism delivering a large home bias generated by the model in Table 3? The driving force of our

[16] The PSID data contain accurate information on wealth holding of households at five-year intervals since 1984. Moreover, the PSID provides weights to map the data to a nationally representative sample. A description of the data is provided in the Appendix.

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Table 3
Aggregate portfolio shares of U.S. investors with liquidity constraints.

	No human capital risk capital risk	Model distribution	Weighted model distribution	Empirical distribution	Weighted empirical distribution
U.S.	25%	95%	75%	75%	61%
Germany	22%	1%	7%	8%	13%
Japan	36%	4%	17%	17%	26%
U.K.	18%	0%	1%	0%	0%
Total	100%	100%	100%	100%	100%

results is that small differences in the correlation of aggregate labor income innovations and market returns, in the presence of short-selling constraints, lead to a gradual international diversification of investors' portfolio as their level of normalized cash-on-hand increases.

In the presence of liquidity constraints, agents cannot borrow to construct an optimally diversified portfolio. Therefore, when their level of liquid wealth to labor income ratio is sufficiently high and they enter the stock market, agents try to minimize the overall wealth risk, investing first in the assets that have the lower degree of correlation with labor income. Only when the ratio of liquid wealth to labor income is sufficiently high, and the labor income risk hedging motive becomes less important relative to the financial risk hedging motive, do agents start diversifying their portfolios. Note that this is exactly the pattern found in the Survey of Consumer Finance data, and reported in Fig. 1.

Since the distribution of liquid wealth to labor income is—in the data as in the model—concentrated in the region of low liquid wealth to labor income ratios, the resulting aggregate portfolio is heavily skewed toward the asset with the lowest correlation with aggregate labor income shocks. Therefore, the human capital hedging motive, once market frictions and idiosyncratic labor income risk are taken into account, is likely to explain a large fraction of the home country bias in several countries.

The above results imply that domestic shocks that lead to a redistribution of total income between capital and labor, therefore lowering the correlation between return on physical and human capital, are likely to skew portfolio holdings toward domestic assets.

In Appendix A.3 we show that very small redistributive shocks (shocks with a variance that is equal to as little as 6%–11% of the output variance), can indeed make labor income innovation more correlated with foreign, rather than domestic, market returns innovation. Many kinds of shocks are expected to have an effect on the income distribution that can rationalize the correlations observed in the data and used in our calibration. Common examples are political business cycles and changes in the bargaining power of unions relative to firms. Among others, the works of Bertola (1993) and Alesina and Rodrik (1994) suggest that changes in the time patterns of capital and labor returns may be the endogenous outcome of majority voting. Santa-Clara and Valkanov (2003) find that in the United States the average excess returns on the stock market are significantly higher under Democratic than Republican presidents. Moreover, if nominal wages and prices have different degrees of stickiness, demand and technological shocks will have redistributive effects on real payoffs to labor and capital. Supportive evidence for redistributive shocks can be found in the empirical literature: Abowd (1989), in a study on wage bargaining in the United States, finds a large and negative correlation between unexpected union wage changes and unexpected changes in the stock value of the firm; Bottazzi et al. (1996), using a VAR approach that imposes block exogeneity across countries (and hence, as discussed in the next sections, is likely to overestimate the benefits of international portfolio diversification), find that the correlations of returns to human capital with domestic market returns is smaller than the one with foreign market returns in 7 out of 10 countries in their study (with an average difference of 0.19); Lustig and Nieuwerburgh (2008)[17] uncover a negative correlation between innovations to human and physical capital returns in the United States; Gali (1999), Rotemberg (2003), and Francis and Ramey (2009) document a negative correlation between labor hours and productivity conditioning on productivity shock.

Note also that the above results have been obtained without considering the exchange rate risk connected with the investment in foreign assets. In the sample period considered, the lower bound on the estimated standard deviation of exchange rates in the three countries considered is about one-third of the standard deviation of market returns. Moreover, the exchange rates show a weakly positive correlation with the stock market of the foreign country and seem to be uncorrelated with the U.S. stock market and with labor income innovations.[18] Therefore, adding exchange rate risk to the model would reduce the Sharpe ratio of foreign assets, making foreign investment less attractive and increasing the degree of home bias.

2.4. Relaxing the borrowing constraints

As a robustness check of the above results we now relax the short-selling constraint restriction. Relaxing this restriction reduces the buffer stock saving need, since short-selling increases the households' ability of smoothing wealth shocks over time via borrowing at the risk free rate (i.e. shorting the risk free asset). In particular, we relax the constraint by allowing the household to borrow (i.e. short-sell) up to a constant fraction of its annual (permanent component of) labor income.

Table 4 computes the aggregate domestic portfolio as in Table 3 but considering a different level of households' borrowing capacity in each of its panels: in Panel A through D, respectively, short-selling is constrained to be no more than 20%, 50%, 100% and 200% of annual labor income.

The table shows that the effect of relaxing the borrowing constraints is non-monotonic. Moderate and intermediate borrowing ability actually increases the degree of home country bias generated by the human capital hedging motive. On the other hand, allowing for extremely large borrowing reduces the degree of home country bias generated by the model.

This non-monotonicity is quite intuitive. With moderate borrowing ability, when the household is sufficiently wealthy to invest in the financial market, it uses its borrowing capacity to leverage and hedge further the human capital risk by skewing holdings toward the domestic stock. As a consequence, the model in this case generates even more home country bias in portfolio holdings than in the baseline specification with no short-selling (reported in Table 3).

When instead the household can borrow large amounts, given the large equity premia, borrowing at the risk free rate to invest in stocks

17 These authors aptly title their paper: "The Returns on Human Capital: Good News on Wall Street is Bad News on Main Street."

18 Hau and Rey (2006) find that, at higher frequencies, higher returns in the home equity market relative to the foreign equity market are associated with a home currency depreciation.

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Table 4
Aggregate portfolio shares of U.S. investors with liquidity constraints—assuming that investors can borrow up to 0.2, 0.5, 1, or 2 times their annual income.

	No human capital risk	Model distribution	Weighted model distribution	Empirical distribution	Weighted empirical distribution
Panel A: Debt limit equal to 0.2 times income					
U.S.	25%	93%	58%	81%	61%
Germany	22%	2%	14%	6%	12%
Japan	36%	5%	27%	13%	26%
U.K.	18%	0%	1%	0%	1%
Total	100%	100%	100%	100%	100%
Panel B: Debt limit equal to 0.5 times income					
U.S.	25%	96%	60%	79%	59%
Germany	22%	1%	13%	5%	13%
Japan	36%	3%	26%	16%	27%
U.K.	18%	0%	1%	0%	1%
Total	100%	100%	100%	100%	100%
Panel C: Debt limit equal to 1 times income					
U.S.	25%	97%	52%	69%	55%
Germany	22%	1%	15%	9%	15%
Japan	36%	2%	30%	22%	28%
U.K.	18%	0%	3%	0%	2%
Total	100%	100%	100%	100%	100%
Panel D: Debt limit equal to 2 times income					
U.S.	25%	89%	48%	57%	50%
Germany	22%	2%	17%	14%	16%
Japan	36%	9%	30%	27%	30%
U.K.	18%	0%	5%	1%	4%
Total	100%	100%	100%	100%	100%

is an attractive investment (for a power utility investor). Since the expected utility from the financial investment is maximised with a well diversified portfolio, a tension between human capital hedging and financial wealth diversification arises. As a consequence, when the household can borrow large amounts relative to the size of its human capital, the financial wealth diversification motive reduces the degree of home country bias in portfolio holdings.

Nevertheless, even with the unrealistically high borrowing capacity considered in panels C and D, the model generates very large home country bias.[19] This is due to the fact that the human capital of the household has a value that is a large multiple of its annual labor income (formally, the present discounted value of all future labor income), while the borrowing capacity, in realistic calibrations, is only a relatively small fraction (or a small multiple in panels C and D) of the current labor income.

3. Measuring factor returns

To assess the role of human capital in determining optimal portfolio choice, and in particular to estimate the correlation between human and physical capital innovations, one needs to study the time series properties of the returns to human and physical capital. This task is complicated by the fact that neither the market value nor the returns to human and (total) physical capital are observable. Nevertheless, total payouts to both factors of production are directly observable from national accounting figures. Moreover, total payouts to the labor force and capital holders can be thought of as the aggregate dividends flows on the unobserved stocks of human and physical capitals. We can therefore use the Campbell and Shiller (1988) methodology to infer the time series properties of unobserved aggregate returns from the observed growth rates of dividends on human and physical capital.

Let P and D be respectively the (unobserved) price and (observed) dividend of an asset. The gross (unobserved) return R is given by the following accounting identity:

$$R_{t+1} := \frac{P_{t+1} + D_{t+1}}{P_t}. \tag{7}$$

Assuming that the price–dividend ratio is stationary, we can log-linearize Eq. (7) around its long-run average to get:

$$r_{t+1} = (1-\rho)\,k + \rho\,(p_{t+1} - d_{t+1}) - (p_t - d_t) + \Delta d_{t+1},$$

where $r_t := log R_t$, $p_t := log P_t$, $d_t := log D_t$, $\Delta d_t := d_t - d_{t-1}$, $\rho := 1/\left(1+\exp(\overline{d-p})\right)$, $\overline{d-p}$ is the long-run average log dividend–price ratio, and k is a constant.

Rearranging the above equation and iterating forward, the log price–dividend ratio can be written (disregarding a constant term) as

$$p_t - d_t = \sum_{\tau=1}^{\infty} \rho^{\tau-1}\left[\Delta d_{t+\tau} - r_{t+\tau}\right] + \lim_{T\to\infty} \rho^T\,(p_{t+T} - d_{t+T}). \tag{8}$$

The equality between the observed log price–dividend ratio, $p_t - d_t$, and future dividend growth rates, $\Delta d_{t+\tau}$, and returns, $r_{t+\tau}$, in Eq. (8) holds for any realization of $\{\Delta l_{t+\tau} - r_{t+\tau}\}_{\tau=1}^{\infty}$ and $p_\infty - dl_\infty$, and hence holds in expectation for any probability measure. This implies that we can take expectations of Eq. (8) under both the

[19] In the Survey of Consumer Finances data (1992–2013), almost all households borrow substantially less than 20% of the household labor income (the case considered in panel A). Moreover, only the calibration in panel A, and to a lesser extent the one in panel B, come close to match the distribution of cash-on-hand in the SCF data, while the calibrations in panels C and D generate too low liquid wealth holdings.

time t and time $t+1$ information sets. Therefore, if we follow Campbell and Shiller (1988) in assuming that the expected one period ahead return is constant, $E_t[r_{t+1}] =: \bar{r}$, and also impose that the transversality condition holds,[20] i.e., $\lim_{T\to\infty} \rho^T(p_{t+T} - d_{t+T}) = 0$, we have that

$$r_{t+1} - \bar{r} = \sum_{\tau=1}^{\infty} \rho^{\tau-1} (E_{t+1} - E_t) [\Delta d_{t+\tau}] \tag{9}$$

where $(E_{t+1} - E_t)[x] := E_{t+1}[x] - E_t[x]$ and E_τ is the rational expectation operator conditional on the information set available up to time τ. Eq. (9) implies that, interpreting the total payouts to labor force and capital holders as the aggregate dividend flows on the unobserved stocks of human and physical capital, we can construct the time series of unobserved returns on human and physical capital as a function of (expected) future growth rates of labor income and capital dividends. The time series of returns constructed in this fashion can then be used to estimate the relevant moments for optimal portfolio choice and human capital hedging.

To make the above approach operational, we need to construct empirical proxies of the expected values in Eq. (9). We perform this task following Campbell and Shiller (1988) and use linear projections generated by a reduced form VAR in a similar fashion as in the seminal work of Baxter and Jermann (1997).

In order to make our results directly comparable with Baxter and Jermann (1997) (and, as discussed below, due to data limitations), we focus on four countries—the United States, the United Kingdom, Germany, and Japan[21] —and the variables included in our VAR are the labor and capital incomes in each of these countries. But, differently from Baxter and Jermann, our empirical procedure allows for common international shocks, comovements, and trends among countries and, as shown in Section 3.1, this difference leads to a sharp difference in results. In particular, the econometric specification used by Baxter and Jermann (1997) relies on the block exogeneity of each country in a VAR framework. Their procedure of estimating a vector error correction model (VECM), a particular case of VAR, separately for each of the four countries is analogous to estimating a VECM for all the countries under the assumption that each country is block exogenous with respect to the other countries. This approach embeds the assumption of low international economic integration. Their VECM, for each country i, takes the form:

$$\begin{bmatrix} \Delta d^i_{L,t+1} \\ \Delta d^i_{K,t+1} \end{bmatrix} = \begin{bmatrix} c^i_L \\ c^i_K \end{bmatrix} + \begin{bmatrix} \psi^i_{LL}(L) & \psi^i_{LK}(L) \\ \psi^i_{KL}(L) & \psi^i_{KK}(L) \end{bmatrix} \begin{bmatrix} \Delta d^i_{L,t} \\ \Delta d^i_{K,t} \end{bmatrix} + \begin{bmatrix} \eta^i_L \\ \eta^i_K \end{bmatrix} \left(d^i_{L,t} - d^i_{K,t} \right) + \begin{bmatrix} \varepsilon^i_L \\ \varepsilon^i_K \end{bmatrix} \tag{10}$$

where $d^i_{L,t}$ denotes the log of labor income, $d^i_{K,t}$ denotes the log of capital income, c^i_L and c^i_K are constant terms, $\Delta d^i_{L,t+1} \equiv d^i_{L,t+1} - d^i_{L,t}$, $\Delta d^i_{K,t+1} \equiv d^i_{K,t+1} - d^i_{K,t}$, and the $\psi_{..}(L)$ terms are polynomials in the lag operator L. The rationale for imposing a cointegration vector of the form $[1, -1]$ is that if labor and capital income are allowed to have independent trends (whether deterministic or stochastic), the labor share of income within a country will reach 1 or 0 with probability 1 as the sample size goes to infinity. Appendix A.4 reports a formal empirical analysis of this assumption.

Eq. (10) can be rewritten in more compact form as:

$$\Delta D^i_{t+1} = C^i + \Psi^i(L)\Delta D^i_t + \Pi^i \left(d^i_{L,t} - d^i_{K,t} \right) + \nu^i_{t+1} \tag{11}$$

where

$$\Delta D^i_{t+1} = \begin{bmatrix} \Delta d^i_{L,t+1} \\ \Delta d^i_{K,t+1} \end{bmatrix}, C^i = \begin{bmatrix} c^i_L \\ c^i_K \end{bmatrix}, \Psi^i(L) = \begin{bmatrix} \psi^i_{LL}(L) & \psi^i_{LK}(L) \\ \psi^i_{KL}(L) & \psi^i_{KK}(L) \end{bmatrix},$$

$$\Pi^i = \begin{bmatrix} \eta^i_L \\ \eta^i_K \end{bmatrix}, \nu^i_{t+1} = \begin{bmatrix} \varepsilon^i_L \\ \varepsilon^i_K \end{bmatrix}.$$

Using this notation and defining ΔD_{t+1} and C as the vectors containing ΔD^i_{t+1} and C^i for each of the four countries considered, the VECM estimated by Baxter and Jermann (1997) can be rewritten as a system of the form

$$\Delta D_{t+1} = C + \begin{bmatrix} \Psi^1(L) & \mathbf{0} & \mathbf{0} & \mathbf{0} \\ \mathbf{0} & \Psi^2(L) & \mathbf{0} & \mathbf{0} \\ \mathbf{0} & \mathbf{0} & \Psi^3(L) & \mathbf{0} \\ \mathbf{0} & \mathbf{0} & \mathbf{0} & \Psi^4(L) \end{bmatrix} \Delta D_t + \begin{bmatrix} \Pi^1 \left(d^1_{L,t} - d^1_{K,t} \right) \\ \Pi^2 \left(d^2_{L,t} - d^2_{K,t} \right) \\ \Pi^3 \left(d^3_{L,t} - d^3_{K,t} \right) \\ \Pi^4 \left(d^4_{L,t} - d^4_{K,t} \right) \end{bmatrix} + \begin{bmatrix} \nu^1_{t+1} \\ \nu^2_{t+1} \\ \nu^3_{t+1} \\ \nu^4_{t+1} \end{bmatrix} \tag{12}$$

where the $\mathbf{0}$ elements are 2×2 matrices of zeros.

There are two important implicit restrictions in Eq. (12). First, the first matrix on the right-hand side of the equation has all the off-diagonal matrices restricted to be zero, i.e., each country is assumed to be block exogenous with respect to the other countries: the first differences of log labor and log capital income of each country are *not* supposed to Granger-cause the first differences of log labor and log capital income in other countries. Second, the cointegration structure in the second term on the right-hand side of Eq. (12) rules out cross-country cointegrations between incomes of the factors of production—that is, it rules out international common trends (e.g., it rules out that capital income in different countries follows the same long-run stochastic trend). Our empirical analysis in the next subsection relaxes both of these restrictions and considers a more general class of VAR models for labor and capital incomes that allow for short- and long-run comovements across countries.

3.1. Empirical evidence: a reappraisal

We estimate Eqs. (9) and (12), as well as alternative VAR specifications, using annual data on labor income and capital income from OECD National Accounts for Germany, Japan, the United Kingdom, and the United States over the period 1960–2012. Our measure of labor income is total employee compensation. This is a less than ideal measure in that it is likely to contain components that are not purely compensation to labor (e.g. the wage bill received by an entrepreneur might contain capital compensation components), and consequently the literature has developed more accurate measures of compensation to labor (see, e.g., Gopinath et al., 2015). Nevertheless, using more accurate measures of wage compensation would require focusing on much shorter time series: depending on the approach, we would lose between 39 and 75 per cent of the time series dimension, and in such a reduced sample it would not be feasible to test for block exogeneity since the number of parameters to be estimated would be too large relative to the number of observations. Similarly, we are constrained to use a relatively small cross-section of countries (that, nevertheless, account more than half of the world equity market capitalization at the end of our sample, and more than 90%

[20] Imposing that the transversality condition holds is less restrictive than it might seem since, even though it rules out intrinsic bubbles as the ones analyzed in Froot and Obstfeld (1991), it does not rule out the presence of mispricings in the asset markets (Campbell and Vuolteenaho, 2004; Brunnermeier and Julliard, 2008).

[21] Baxter and Jermann (1997) focus on this sample since they estimate the cumulative share of these four countries in the world portfolio to be around 93%.

Table 5
Testing block exogeneity.

	Test statistic
Likelihood ratio	83.7 (0.001)
Wald	91.8 (0.000)
Lagrange multiplier	71.2 (0.016)

Note: Tests for the block exogeneity restriction for the VECM in Eq. (12). p-value of not rejecting the null hypothesis is in brackets.

of the world capitalization at the beginning of the sample), since expanding to more countries would not only substantially increase the number of parameters to be estimated in each equation of the VAR system,[22] but also shorten the available time series since the countries we consider are exactly the ones with the longest available history of wage bill data. This data limitation, nevertheless, has the advantage of making our results directly comparable to the previous literature and, in particular, to Baxter and Jermann (1997) since we use exactly the same set of countries and definition of the wage bill.

Our baseline measure of capital income is GDP at factor cost minus employee compensation. A detailed description of the dataset is given in Appendix A.1.

3.1.1. *Model selection*

The restrictions imposed in Eq. (12) by Baxter and Jermann (1997) can be formally tested against more general specifications that allow for international comovements in the payoffs to the factor of production. We start by assessing whether the block exogeneity assumption is supported by the data. Table 5 reports frequentist tests of the null hypothesis of block exogeneity in Eq. (12). Both restricted and unrestricted specifications are estimated with only one lag (as in Baxter and Jermann, 1997), and we maintain the hypothesis of cointegration relationships only within the countries as in Eq. (12). As stressed by the p-values reported under the test statistics, the null hypothesis of block exogeneity is rejected at any standard confidence level. That is, the data suggest that there exist statistically significant cross-country links between the compensations of the factors of production.

Next, we want to relax the hypothesis of only within countries cointegration relationships. That is, we want to allow for cross-country common trends across variables. Since any possible VECM representation of the data would have a one-to-one mapping to a corresponding VAR model in levels of log labor and capital income, we do so by considering this last class of models. In comparing VARs in levels against other specifications, we need to be careful about the unit roots in the labor and capital income series. [23] As a consequence, we perform model comparison employing Bayesian posterior probabilities (see, e.g., Gelman et al., 1995) since this approach is robust to non-stationarity in the data. In particular, for each model j considered, we compute the Bayes factor

$$BF_j = \int_{\Theta_j} g_j(\theta) p_j(X \mid \theta) d\theta \tag{13}$$

where $p_j(X \mid \theta)$ is the likelihood function of the data, X, θ is a vector of parameters belonging to the space Θ_j, and $g_j(\theta)$ is a prior on the distribution of the parameters of the jth model. Using the Laplace method (see, e.g., Schervish, 1995) the Bayes factor can be approximated as

$$BF_j \cong g_j(\hat{\theta}_j) p(X \mid \hat{\theta}_j) (2\pi)^{\frac{m}{2}} \left| \hat{\Sigma}_{\theta j} \right|^{\frac{1}{2}} \tag{14}$$

where $p_j(X \mid \hat{\theta}_j)$ is the likelihood of the j model evaluated at its peak $\hat{\theta}_j$, m is the dimension of Θ_j, and $\hat{\Sigma}_{\theta_j}$ is the observed information matrix. Note that the above approximation is accurate even in the presence of unit roots (see, e.g., Kim, 1994).

With the Bayes factors at hand, the posterior probability of each model j is computed as

$$PO_j = \frac{p_j BF_j}{\sum_i p_i BF_i} \tag{15}$$

where p_j is the prior probability of the jth model.

Table 6 reports the logs of the Bayes factor and the posterior probabilities defined by Eqs. (14) and (15) for a large set of models, under the assumption of flat priors and equal prior probability for each model. The models considered are as follows: *i*) vector error correction models—with (row 1) and without (row 2) block exogeneity restrictions—in which, as in Baxter and Jermann (1997), the only cointegrations allowed are within country and the fixed cointegration vector has the form $[1, -1]$; *ii*) VARs in levels with the block exogeneity restriction which relax the assumption of having a cointegration vector of the form $[1, -1]$ (rows 3 and 4); *iii*) VARs in first differences (with, row 5, and without, row 6, block exogeneity) which rule out any form of cointegration among variables; and *iv*) unrestricted VARs in levels (rows 7 and 8) that allow for international comovements and arbitrary cross-countries—as well as within country—cointegration relationships. The maximum number of lags considered for each specification is naturally restricted by the sample size at hand, but nevertheless corresponds to the one chosen by Akaike and Bayesian information criteria.

The econometric model considered in row 1 of Table 6 corresponds to the original Baxter and Jermann (1997) specification. The second row shows that relaxing the block exogeneity assumption leads to a dramatic increase in the log Bayes factor ($\log BF_i$). This increase is so large that if the models in the first two rows were the only ones considered, we would assign a posterior probability of about one to the specification that—unlike Baxter and Jermann—

Table 6
Log Bayes factors and posterior probabilities.

Row	Specification	log BF_j	PO_j
(1)	VECM with block exogeneity, domestic cointegration, one lag	724.35	$1.26e-51$
(2)	VECM without block exogeneity, domestic cointegration, one lag	790.15	$4.76e-23$
(3)	VAR in levels with block exogeneity, one lag	701.22	$1.14e-61$
(4)	VAR in levels with block exogeneity, two lags	725.06	$2.63e-51$
(5)	VAR in first-differences with block exogeneity, one lag	717.51	$1.38e-54$
(6)	VAR in first-differences without block exogeneity, one lag	781.58	$9.02e-27$
(7)	VAR in levels without block exogeneity, one lag	769.23	$3.91e-32$
(8)	VAR in levels without block exogeneity, two lags	841.55	1

Note: Logs of Bayes factors and posterior probabilities. The posterior probabilities do not sum up to 1 because of rounding error.

[22] E.g., adding even only one country to our setting would imply, in the best specification case of Table 6 below estimating an additional 23 parameters—i.e. a dramatic reduction in degrees of freedom given the size of the time series dimension of the data (52 years of annual observations).

[23] The unit root hypothesis cannot be rejected for all the data series considered.

allows for international comovements among variables. That is, Bayesian testing confirms the strong rejection of the block exogeneity assumptions delivered by frequentist testing presented in Table 5. The models considered in the third and fourth rows maintain the block exogeneity assumption but, by considering VARs in levels, do not restrict the within country cointegration vector to take the form $[-1, 1]$. The Bayes factors of these specifications are of similar magnitude to the one in the first row, but much smaller than the one in the second row, providing additional evidence of a strong rejection of the block exogeneity assumption. The VARs in first differences with and without block exogeneity in rows 5 and 6 are relevant because they impose the restriction of no cointegration among variables. Since the Bayes factor in row 5 is smaller than the one in row 1, and the one in row 6 is smaller than the one in row 2, the data provide supporting evidence for within country cointegration in the payoffs to physical and human capital. Nevertheless, in Appendix A.4, we report a detailed frequentist analysis of within country cointegration and find mixed evidence in support of this hypothesis: the $[-1, 1]$ cointegration vector is always rejected except for Japan, while relaxing this parameter restriction the results vary from country to country and with the lag length considered.

Finally, the specification in rows 7 and 8 of Table 6 are unrestricted VARs in levels with one and two lags, respectively. These specifications allow for arbitrary cointegration within and, most importantly, across countries—that is, the variables are allowed to show both short- and long-run systematic comovements across countries. The specification with two lags (which can be mapped into a VECM) in row 8 delivers a Bayes factor that is substantially higher compared to all the other models considered. This large Bayes factor maps into a posterior probability (PO_j in the second column of Table 6) that is numerically indistinguishable from 1. That is, the data provide strong evidence of both short- and long-run cross-country comovements in the payoffs to human and physical capital, implying that the econometric model of Baxter and Jermann (1997)—that rules out both of these channels—is misspecified.

To test the robustness of the above results, we use numerical integration of Eq. (13) (we used an importance sampling approach based on the asymptotic Normal-inverse-Wishart shape of the posterior to perform this task) to get alternative estimates of the Bayes factors and posterior probabilities in Table 6. We also experimented with non-flat priors over the parameters space. In both cases, the results are in line with the ones in Table 6.

Overall, the results of this subsection imply that to accurately measure returns to human and physical capital, and their implications for international portfolio diversification, we should use an econometric specification that, differently from the ones used in the previous literature, allows for both short- and long-run international comovements in the payoffs to production factors.

3.1.2. The correlation of human and physical capital returns

In order to estimate factor returns using Eq. (9) and the selected VAR model, we calibrate the parameters ρ to 0.957 for both capital and labor income. This corresponds to assuming that the mean dividend–price ratio of labor income and capital income are identical and equal to 4.5% as in Baxter and Jermann (1997). Moreover, note that finiteness of the empirical estimate of the right-hand side of Eq. (9) is guaranteed if ρ times the largest eigenvalue of the companion matrix of the selected VAR model is within the unit circle. This condition is satisfied by our choice of ρ.

Table 7 reports the correlations between returns on capital and labor computed using Eq. (9) and the estimations of expected Δd's by the VAR in levels specification with two lags.

The correlations are both qualitatively and quantitatively different from the ones derived by Baxter and Jermann (1997). The within countries correlations seem to be somewhat lower compared to Baxter and Jermann (1997): their estimates cover the range

Table 7
Correlation of factor returns.

	r_K^G	r_L^J	r_K^J	r_L^{UK}	r_K^{UK}	r_L^{USA}	r_K^{USA}
r_L^G	0.761 [0.3,0.94]	0.701 [0.2,0.94]	0.828 [0.59,0.97]	0.727 [0.26,0.95]	0.747 [0.28,0.94]	0.847 [0.55,0.97]	0.808 [0.42,0.96]
r_K^G		0.144 [−0.55,0.73]	0.725 [0.14,0.95]	0.869 [0.55,0.99]	0.986 [0.95,1]	0.933 [0.76,0.99]	0.977 [0.92,1]
r_L^J			0.666 [0.15,0.93]	0.155 [−0.53,0.77]	0.170 [−0.52,0.74]	0.311 [−0.4,0.83]	0.239 [−0.48,0.78]
r_K^J				0.524 [−0.11,0.93]	0.751 [0.2,0.96]	0.738 [0.27,0.97]	0.739 [0.22,0.96]
r_L^{UK}					0.861 [0.55,0.99]	0.945 [0.8,0.99]	0.918 [0.73,0.99]
r_K^{UK}						0.933 [0.77,0.99]	0.982 [0.94,1]
r_L^{USA}							0.964 [0.87,0.99]

Note: Correlations of human capital returns with physical capital returns. Factor returns are estimated using a VAR specification in levels with two lags. We report in brackets the 95% confidence interval constructed using sampling-with-replacement raw residuals bootstrap based on 10,000 replications.

[0.78, 0.99], while our estimates have a maximum of 0.96 and a minimum of 0.67 in Japan.[24]

The between countries correlations appear to be higher compared to Baxter and Jermann (1997): their maximum correlation between returns on capital is 0.43 (U.S.–Germany), the maximum correlation between returns on labor is 0.35 (U.S.–Germany), the maximum correlation between domestic labor returns and foreign capital returns is 0.40 (Germany–U.S.).

In our estimation, the between countries correlations for both r^L and r^K are much higher (with the exception of Japan, where the correlations between labor returns and foreign returns on both capital and labor are generally lower). The correlations between returns on capital, for example, cover the range [0.73, 0.98]. Moreover, for all countries but Japan, the correlations between domestic returns on labor and foreign returns on capital are similar to the correlation between domestic returns on labor and capital. These results suggest the presence of productivity shocks effective at the international levels.

The factor returns that we obtain by applying Eq. (9) are generated regressors (Oxley and McAleer, 1993). To account for this, we compute the standard errors of the corresponding correlation matrix using a bootstrap approach to statistical inference (see, e.g., Efron and Tibshirani, 1993). More specifically, we apply a sampling-with-replacement raw residuals bootstrap scheme with 10,000 repetitions. Interestingly, as shown in Table 7, we find large empirical confidence intervals. For Japan, the confidence intervals indicate that the correlation between factor returns may be either positive or negative. In other words, we document the existence of substantial statistical uncertainty on measuring returns to the aggregate capital stock.

The differences between our point estimates and the results of Baxter and Jermann (1997) are mostly driven by the relaxation of the block exogeneity assumption, which turns out to be strongly rejected by the data. Baxter and Jermann (1997) fit a model where they restrict the countries not to be economically and technologically integrated. As an outcome, the level of between countries correlation is underestimated and the within country correlation is overestimated (i.e., the countries appear not to be integrated). Once this restriction is removed, the effective degree of technological and economic integration becomes evident. This high degree of economic integration implies fewer opportunities to hedge the human capital risk investing in foreign marketable assets.

[24] Bottazzi et al. (1996) estimate a negative correlation between wage rate and domestic profit rate. Their estimations are much more in line with the within country correlations reported in Table 3 than with the ones reported by Baxter and Jermann (1997).

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For comparability, in Table A5 of Appendix A.6, we replicate the optimal portfolio implied by the complete spanning approach of Baxter and Jermann (1997). That is, we replicate the authors' main result, but after correcting their VAR misspecification (i.e. we use the VAR specification in Row (8) of Table 6). Raw point estimates indicate that the authors' original claim, i.e. that the home country bias is generally worsened by the human capital hedging motive, is not generally supported by the data: the table shows that, for some countries, we obtain exactly the opposite effect. Moreover, confidence intervals show that there is substantial uncertainty about optimal portfolios constructed with this approach, and that human capital hedging can potentially generate large home country bias.[25]

3.1.3. The correlation of human capital and stock market returns

Estimating returns to physical capital using the approach in Eq. (9) might be inappropriate for evaluating international portfolio diversification, since *i*) only a subset of the claims to capital compensation in a country are tradable internationally with relatively little frictions—i.e., the ones of publicly traded companies, and *ii*) due to tax advantages, the compensation to capital elicited from national accounts tends to include de facto components of human capital compensation (e.g., for family owned and individual firms). As a consequence, a more appropriate measure of the correlation between human and physical capital compensations can be constructed replacing the VAR based estimates of returns to capital with the returns on broad stock market indices. In this subsection, we restrict the set of assets available to hedge human capital to include only publicly traded stocks, since we consider this as being the relevant case for most households. The innovations to human capital compensation are estimated as in the previous subsections.

Table 8 reports the correlations between the returns to human capital and returns to the stock market. Since the correlation between U.S. human capital innovations and stock returns plays a key role in the model calibration presented in Section 2, for the United States we use three different stock indices: the Fama and French (1992) benchmark market return,[26] the S&P 500 index, and the Dow Jones Industrial index.

Overall, the table suggests that domestic returns to human capital are in general more correlated with foreign stock markets: for all countries considered, the asset with the highest correlation with human capital returns is always a foreign asset. Moreover, in the United States in particular, returns to human capital have the lowest correlation with the domestic, rather than the foreign, stock market index. However, the estimated correlations tend to be quantitatively small (this is consistent with the findings of Fama and Schwert (1977)), and the uncertainty attached to the estimation of human capital delivers large confidence intervals (in particular, only one of the estimated correlation coefficients is different from zero at the 95% confidence level). Nevertheless, as shown in Table A4 of Appendix A.5, despite the large confidence intervals, the pattern of lower correlation of human capital with domestic, rather than foreign, returns, is quite robust to alternative construction of the data.

The small correlations in the tables have two important implications. First, when focusing on tradable claims to physical capital, the assumption of complete spanning for human capital return used in the previous literature seems to be unsupported by the data. Second, if one were to use a value weighted approach for the determination of optimal portfolios as in Baxter and Jermann (1997), the effect of human capital hedging would be very small, but it would tend to skew holdings in favor of domestic assets, as shown in Table A5 in Appendix A.6.

Table 8
Correlation of factor returns using stock market data.

	r_L^G	r_L^J	r_L^{UK}	r_L^{USA}
r_K^G	0.118 [−0.08,0.24]	0.138 [−0.8,0.24]	0.110 [−0.07,0.26]	0.145 [−0.04,0.27]
r_K^J	0.122 [−0.05,0.21]	0.071 [−0.09,0.19]	0.145 [−0.02,0.23]	0.194 [0.03,0.26]
r_K^{UK}	0.130 [−0.09,0.26]	0.105 [−0.12,0.23]	0.101 [−0.10,0.27]	0.149 [−0.06,0.29]
r_{FF}^{USA}	0.000 [−0.14,0.11]	−0.088 [−0.20,0.04]	0.031 [−0.10,0.15]	0.041 [−0.10,0.14]
r_{DJ}^{USA}	−0.017 [−0.18,0.12]	−0.147 [−0.27,0.01]	0.035 [−0.07,0.21]	0.006 [−0.10,0.18]
r_{SP}^{USA}	0.026 [−0.11,0.12]	−0.076 [−0.19,0.05]	0.075 [−0.05,0.17]	0.070 [−0.06,0.16]

Note: Correlations of human capital returns with physical capital returns. Returns to human capital are estimated using a VAR specification in levels with two lags. Returns on physical capital are computed using stock market data (CDAX for Germany; Nikko Securities Composite for Japan; FTSE All-Share Index for United Kingdom; Fama and French (FF), S&P 500 Total Return Index (SP), and Dow Jones Industrials Total Return Index (DJ) for the United States). We report in brackets the 95% confidence interval constructed using sampling-with-replacement raw residuals bootstrap based on 10,000 replications.

For robustness, we have also estimated the correlations in Table 8 using two subsamples of equal length (pre and post 1987). The correlations in the two subsample are not statistically different from the full sample ones, albeit smaller in the second subsample. More importantly, the subsample estimates highlight the same pattern as in Table 8: labor income innovations tend to be more correlated with foreign stock markets than domestic ones. [27]

Nevertheless, as we have shown in Section 2 above in the presence of borrowing constraints and both aggregate and idiosyncratic human capital risk, even the above small correlations can have a very large impact on optimal portfolio decisions.

4. Conclusion

This paper shows that human capital risk can help rationalize the home country bias in equity holdings at both the aggregate and household levels. First, we show that the theoretical intuition that short positions in domestic physical assets are a good hedge for human capital risk is a very fragile one, as very small redistributive shocks—e.g., with a variance of a mere 6% of GDP variance—are enough to reverse this intuition. Moreover, we find that the presence of this type of shocks is supported by the data.

Second, we show that the commonly used approach of estimating country-specific VARs to compute returns to human and physical capital is rejected by the data and delivers mechanically biased estimates of the hedging benefits of shorting the domestic capital. Most importantly, we show that this misspecification largely drives the findings of Baxter and Jermann (1997)—i.e., the result that the home bias puzzle is unequivocally worse than we think once we consider human capital risk, is the outcome of an econometric misspecification strongly rejected by the data.

Moreover, we show that when returns to physical capital are measured using broad stock market indexes, human capital return innovations tend to be more correlated with foreign rather than domestic stocks. Nevertheless, these correlations are small and, consequently, in a frictionless complete market setting, have very little impact on optimal portfolios.

[25] This last finding also reconciles the discrepancy between the optimal portfolios estimated in Baxter and Jermann (1997), and the ones constructed by Bottazzi et al. (1996).

[26] This is a very broad index that covers includes all NYSE, AMEX, and NASDAQ firms and is available from Kenneth French's website: http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/

[27] For instance, US labor income return innovations have a correlation with foreign stock market returns of 0.19–0.26 in the first sub sample, and 0.02–0.18 in the second subsample, while the correlation of US labor income innovation and US market returns is about 0.04–0.10 in the first subsample and between −0.01 and 0.01 in the second subsample.

Fourth, calibrating a buffer stock saving model consistent with both micro and macro labor income dynamics—hence taking into account that individual labor income uncertainty is substantially larger than the aggregate one— as well as stock market data, we show that a large home country bias arises as an equilibrium result. This is due to the fact that household labor income risk is about one order of magnitude larger than aggregate labor income risk and, in the presence of liquidity constraints, optimal hedging becomes heavily skewed toward the asset whose innovations have the lowest correlation with the labor income innovations—the domestic asset.

Moreover, our heterogenous agents model implies that, at the household level, the degree of home country bias should increase, and portfolio diversification decrease, in the labor income to financial wealth ratio—and these are exactly the novel empirical stylized facts that we uncover in the 1992–2013 U.S. Survey of Consumer Finances data.

A. Appendix

A.1. Data description

In order to allow for a direct comparison of our results with the results of Baxter and Jermann (1997), we collect annual data on labor income and capital income from the OECD National Accounts for Japan, the United Kingdom, and the United States. The data for the period from 1960 to 2003 are available through ESDS International (http://www.esds.ac.uk). More recent data (i.e., data for the period from 1970 to 2012) are available directly from the OECD's iLibrary (http://stats.oecd.org/). Our measure of labor income is total employee compensation paid by resident producers (Table 3, number 1). Our measure of capital income is GDP at factor cost (Table 3, number 18) minus employee compensation.

This leaves us with two overlapping datasets that need to be spliced to cover the full sample period from 1960 to 2012. Unfortunately, the calculation standards for GDP have changed over time and are thus slightly different for the two datasets. Therefore, we cannot simply append more recent data to the older dataset, so we use the ratio of GDP for one dataset in the year in which we splice the data to GDP for the other dataset as a conversion rate. In our baseline specification, we augment the dataset from 1960 until 2003 with more recent data that is adjusted using the GDP ratio for the year 2003 as a conversion rate. Appendix A.5 discusses the robustness of our results with respect to the construction of the labor and capital income data series.

Finally, the OECD National Accounts database for Germany refers to data for the whole of Germany prior to the German reunification, using the ratio of GDP for West Germany to GDP for Germany as a whole in 1991 as a conversion rate. Hence, to better take into account the legal barriers to international investments faced by East Germans prior to 1990, we construct labor and capital income series for West Germany only by employing data from the German Statistical Office (http://www.vgrdl.de/Arbeitskreis_VGR).

The summary statistics of the labor and capital income series are reported in Table A1. The sample averages of the labor shares

Table A1
Summary statistics.

		Labor income	Capital income	Labor share
Germany	Mean	190943	170547	0.530
(millions of euros)	Std. deviation	62059	63631	0.029
Japan	Mean	154322	121685	0.508
(billions of yen)	Std. deviation	62670	46028	0.047
U.K.	Mean	70719	53747	0.573
(millions of pounds)	Std. deviation	19304	18160	0.025
U.S.	Mean	150265	109329	0.580
(millions of dollars)	Std. deviation	43485	34474	0.011

Note: The sample period is 1960–2012.

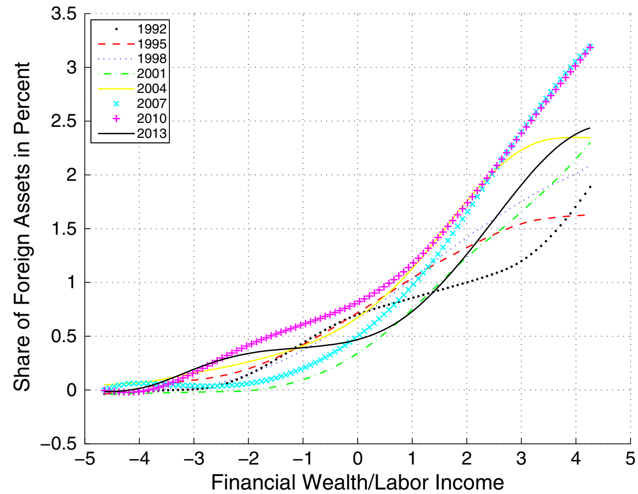

Fig. A1. Locally weighted regressions of the share of foreign assets on the logarithm of financial wealth divided by a household's labor income for different years of the U.S. Survey of Consumer Finances.

of income reported in the table are the values for α_j used in Eq. (A2).

We retrieve the following stock market data from the Global Financial Data (https://www.globalfinancialdata.com): CDAX Total Return Index for Germany, Nikko Securities Composite Total Return for Japan, FTSE All-Share Return Index for the United Kingdom, and S&P 500 Total Return Index and Dow Jones Industrials Total Return Index (DJ) for the United States. The Fama and French benchmark return for the United States (http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/data_library.html) is given by the sum of the Fama and French excess market return and the risk-free rate.

In order to construct real per capita returns, we collect population data and GDP figures deflator from the International Financial Statistics service of the International Monetary Fund (available through http://ukdataservice.ac.uk/).

Survey of Consumer Finances data and the corresponding codebooks are available from the Federal Reserve (http://www.federalreserve.gov/econresdata/scf/scfindex.htm). Data was collected for 8 different surveys from 1992 to 2013. The questionnaires are slightly different for different years. The questions regarding foreign stock holdings, however, remain the same over time. Fig. A1 shows the results of locally weighted regressions of the share of foreign assets on the logarithm of normalized financial wealth, i.e. financial wealth divided by a household's labor income.

A.2. Estimation of the aggregate labor income process

In order to model the U.S. labor income process, we experimented with several specifications in the ARIMA class and performed the standard set of Box–Jenkins selection procedures. In particular, among the models considered, MA(2) and ARMA(1,1) processes fit well to first differences of log labor income. Since these specifications deliver similar results, we henceforth restrict attention to the ARIMA(0,1,2) specification for log income since it simplifies the exposition and it has previously been used in the literature in similar contexts (see, e.g., Davis and Willen, 2000; MaCurdy, 1982). Thus,

Table A2
Estimated labor income process.

$\hat{\mu}_y$	$\hat{\vartheta}_1$	$\hat{\vartheta}_2$	St. error of $\hat{\varepsilon}$
0.0188	0.4475	0.0937	0.0214
(0.0045)	(0.1501)	(0.1556)	

Note: Newey–West standard errors reported in brackets.

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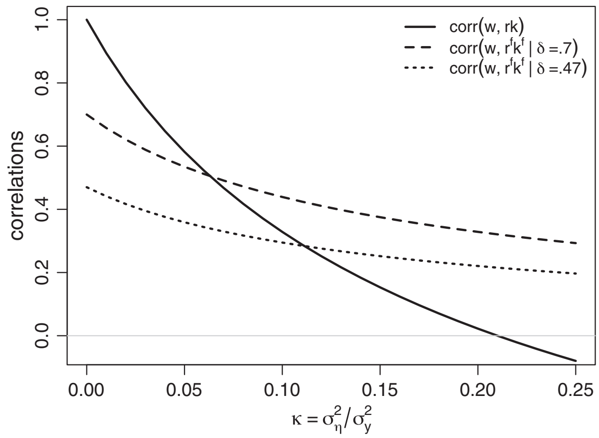

Fig. A2. Correlations of domestic wage and: domestic (solid line) and foreign (dashed and dotted lines) capital compensation for different levels of international GDP correlation (δ).

the fitted earning specification is the MA(2) process Eq. (2). The estimated coefficients are reported in Table A2 below.

A.3. Rent shifting shocks and the home country bias

The empirical results obtained in Sections 2 and 3 stand in sharp contrast with the canonical intuition that, in a competitive and frictionless world with constant labor and capital income shares, payoffs to domestic capital and labor are perfectly correlated. Therefore, if outputs are not perfectly correlated across countries, human capital hedging should skew domestic portfolios toward foreign assets. In this section, we present a stylized calibration exercise and show that, once the degree of international economic integration observed in the data is accounted for, very small redistributive shocks are sufficient to reverse this conclusion: that is, we show that the canonical intuition is theoretically fragile.

Consider a simple two-country setting. The domestic output is produced by a constant return to scale Cobb–Douglas production function with constant elasticity of output to capital and labor inputs—denoted respectively as α and $1-\alpha$. Assume also that there are additive redistributive shocks (η) to factor remuneration—that is labor and capital compensation are given, respectively, by

$$w = (1-\alpha)\,y + \eta$$
$$rk = \alpha y - \eta$$

where y denotes domestic output per capita, w is the domestic wage rate, r is the return on capital, k is the per capita stock of capital, and for simplicity of exposition the redistributive shock η is assumed to be a mean zero iid stochastic process.

Without loss of generality, assume that $\sigma_\eta^2 = \kappa\sigma_y^2$, i.e. κ measures the ratio between the variance of the redistributive shock (σ_η^2) and the per capita output variance (σ_y^2). Assume also that the foreign economy is characterized by the same structure as the domestic economy, with $\sigma_y^2 = \sigma_{y^f}^2$, $\kappa^f = \kappa$, and $\alpha^f = \alpha$, where the superscript f denotes quantities in the foreign country. Denote with $\delta = corr(y, y^f)$ the correlation between domestic and foreign output (y^f). It then follows that

$$corr\,(w, rk) = \frac{(1-\alpha)\,\alpha - \kappa}{\sqrt{(\alpha^2+\kappa)\left[(1-\alpha)^2+\kappa\right]}}$$
$$corr\left(w, r^f k^f\right) = \frac{(1-\alpha)\,\alpha}{\sqrt{(\alpha^2+\kappa)\left[(1-\alpha)^2+\kappa\right]}}\delta$$

where $r^f k^f$ is the payout to capital invested in the foreign country, implying that

$$sign\left(corr\,(w, rk) - corr\left(w, r^f k^f\right)\right) = sign\,((1-\alpha)\,\alpha\,(1-\delta) - \kappa)\,.$$

This simple observations implies that for sufficiently high volatility of the redistributive shock relative to output volatility—i.e., for sufficiently high κ—the domestic asset becomes a better hedge for human capital than the foreign asset. Moreover, this effect becomes stronger when the correlation between domestic and foreign output (δ) increases. The relevant question is whether unrealistically large redistributive shocks and cross-country output correlations are needed for this effect to arise.

Calibrating the capital share of output as $\alpha = .3$, Fig. A2 plots the correlation between domestic labor income and domestic capital income (solid line), and the correlation of domestic labor income with foreign capital income (for different values of δ in the dashed and dotted lines) as a function of κ (the ratio of the redistributive shocks variance to output variance). The two values of δ considered are the minimum and maximum correlations between domestic output and world output reported in (Obsfeld and Rogoff, 1996, page 291) for Germany, the United Kingdom, the United States, and Japan.[28] The figure shows that a redistributive shock with a variance that is a mere 6% (11%) of output variance is enough to make the domestic asset a better hedge for human capital than the foreign asset when the correlation between domestic and foreign output, δ, is 0.7 (0.46). That is, relatively small redistributive shocks are sufficient to revert the intuition that, in a Cobb–Douglas world, hedging labor income risk should skew domestic portfolios toward domestic assets. A natural question is whether redistributive shocks of this magnitude are realistic. A simple way to gauge the magnitude of κ is to compute the ratio of *i*) the variance of the residual of a linear regression of de-trended labor income on de-trended output, to *ii*) the variance of de-trended output. Such an estimate delivers a value for κ in the four countries considered in our empirical analysis (the United States, the United Kingdom, Japan, and Germany during the 1960–2012 period)[29] that ranges from 0.06 to 0.48, when using Hodrick and Prescott (1997) de-trending, and from 0.09 to 0.18 when using linear de-trending. That is, these estimates are in the range needed to skew portfolio holdings toward domestic assets.[30] Moreover, in the presence of rent shifting shocks, we would expect to observe time variation in the labor share of income. Fig. A3 plots the labor income share in total output in the four countries considered in our empirical analysis. The figure shows substantial time variation in the share of GDP received by the labor force.

Note also that the patterns in Fig. A3, from the late seventies onward, show a negative trend in the labor share consistent with the findings of Karabarbounis and Neiman (2013). These author study a large cross-section of countries and document that the global labor share has significantly declined since the early 1980s, with the decline occurring within the large majority of countries and industries, and with a 5 percentage point global decline in the labor share since the late seventies.

[28] Using an extended sample that runs until 2012, we find much higher correlations, hence the calibration reported in the figure is conservative.

[29] A detailed description of the data used throughout the paper is reported in Appendix A.1.

[30] We find very similar results if we gauge the magnitude of κ by computing the ratio of *i*) the variance of the residual of a linear regression of de-trended capital income on de-trended output, to *ii*) the variance of de-trend output.

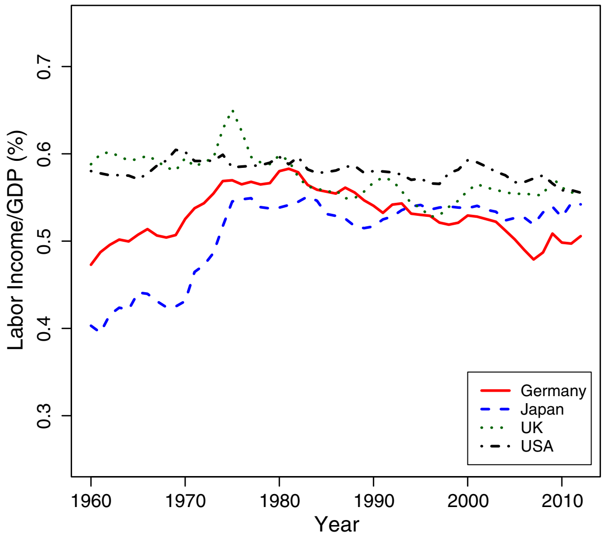

Fig. A3. Labor share of income in Germany, Japan, UK and USA

A.4. Cointegration analysis

In this Appendix, we provide a battery of econometric tests on the time series properties of labor and capital income in the four countries in our sample.

Since both the Augmented Dickey–Fuller (Dickey and Fuller, 1981) and the Phillips–Perron (Phillips and Perron, 1988) tests cannot reject the null hypothesis of unit roots in labor and capital income (results not reported but available upon request), we investigate whether the logarithm of labor income shares a common stochastic trend with the logarithm of capital income, thus implying the existence of a pairwise cointegration relationship within each country.

Table A3 presents econometric evidence based on the Johansen cointegration test (Johansen, 1995), in which we allow for a linear deterministic trend in the data and an intercept (but not a trend) in the cointegration vector. For each country, the table reports both the trace statistics and the maximum eigenvalue statistics for different lags specifications of the first differenced terms. The last column of Table A3 tests the restriction that the log ratio of labor income to capital income is stationary, i.e., the log of labor income and the log of capital income are cointegrated with cointegrating vector $[1, -1]$.

The null hypothesis of no cointegration between labor and capital income cannot be generally rejected at standard significance levels for all countries but Japan. Hence, the long-run restriction suggested by Baxter and Jermann (1997) of pairwise cointegration within countries is not supported by the data unless the number of lags in the test equation is zero. The sensitivity of the result to the lag specification can be due to the low power of the cointegration test, stemming perhaps from a small sample size.

In Fig. A1 we plot the log ratio of labor income to capital income in the four countries in our sample. Simple visual inspection suggests that the behavior of the log ratio of labor income to capital income is not stationary. Indeed, as reported in the last column of Table A3, a formal testing procedure based on the Johansen cointegrating test strongly rejects the cointegration restriction for all countries.

We acknowledge that if labor and capital income are allowed to have independent trends, then the ratio of labor income to capital income either will grow without bound or approach zero asymptotically. Therefore, labor's share will approach one or zero with probability one. However, it is important to realize that the cointegration tests are simply rejecting the hypothesis about the existence of a pairwise cointegration relationship within each country, but are completely silent about the existence of a long-run equilibrium relationship across countries. Moreover, since in this paper we do our econometric analysis in levels, we allow for implicit cointegrating relationships in the data.

Table A3
Johansen cointegration test.

Country	Number of lags	Trace	Max Eigenvalue	Likelihood ratio
Germany	0	20.22 (0.009)	19.80 (0.006)	15.32 (0.002)
	1	11.377 (0.190)	9.504 (0.272)	7.196 (0.007)
	2	10.159 (0.302)	8.59 (0.370)	6.39 (0.012)
	3	10.48 (0.270)	8.34 (0.396)	5.49 (0.019)
Japan	0	54.96 (0.001)	51.84 (0.001)	4.36 (0.037)
	1	27.42 (0.001)	22.86 (0.002)	3.70 (0.055)
	2	17.40 (0.026)	12.06 (0.109)	4.37 (0.037)
	3	30.04 (0.001)	26.04 (0.01)	19.83 (0.000)
United Kingdom	0	18.63 (0.016)	17.46 (0.015)	13.58 (0.000)
	1	12.41 (0.0139)	9.91 (0.229)	4.98 (0.026)
	2	11.30 (0.194)	8.39 (0.391)	5.23 (0.022)
	3	12.30 (0.143)	10.48 (0.183)	7.62 (0.006)
United States	0	16.53 (0.035)	15.85 (0.028)	12.96 (0.000)
	1	12.28 (0.144)	10.61 (0.175)	8.76 (0.003)
	2	13.43 (0.100)	8.88 (0.339)	3.87 (0.049)
	3	15.53 (0.049)	10.53 (0.180)	5.51 (0.019)

Note: Johansen cointegration test. The null hypothesis is no pairwise cointegration between labor income and capital income within each country. We report the trace and maximum eigenvalue test statistics for different lag specifications of the first differenced terms. The last column tests the restriction that the log ratio of labor income to capital income is stationary, i.e., the log of labor income and the log of capital income are cointegrated with cointegrating vector $[1, -1]$. We report the p-value of the test statistics in parentheses.

A.5. Robustness of factor returns correlations

This section tests whether the correlation of the factor returns presented in Table 8 are sensitive to the way we construct our data series on labor income. As mentioned in Appendix A.1, our final data of labor and capital income stems from two different datasets that are then combined. An immediate concern is whether the results of this paper are robust to the different possible ways of combining the two datasets. In particular, one could think of four different ways of combining them: 1) augment the 1960–2003 dataset from 2003 onwards (baseline specification); 2) augment the 1960–2003 dataset from 1970 onwards; 3) augment the 1970–2012 dataset from 2003 backwards; 4) augment the 1970–2012 dataset from 1970 backwards. We opt for the first alternative for several reasons. First, we choose to augment the 1960–2003 dataset since this maximizes the overlap with the data Baxter and Jermann (1997) used in their analysis. Second, we decide to splice the datasets in the year 2003 because, at that time, Germany had already reunified, and consequently the data definitions are consistent across countries. Nevertheless, we explore the robustness of our results by comparing them for the four different alternatives.

In Table A4, we recalculate Table 8 for the four alternatives. The results for alternatives 1) to 4) are reported clockwise, with the top left part being identical to Table A5—our baseline specification. The results suggest that, keeping the year in which we splice fixed, there is essentially no difference in the correlations whether we extend the 1960–2003 or the 1970–2012 dataset. This can be seen from

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Table A4
Correlation of factor returns using stock market data.

1) Augment 1960–2003 from 2003 onwards

	r_L^G	r_L^J	r_L^{UK}	r_L^{USA}
r_K^G	0.118 [−0.08,0.24]	0.138 [−0.8,0.24]	0.110 [−0.07,0.26]	0.145 [−0.04,0.27]
r_K^J	0.122 [−0.05,0.21]	0.071 [−0.09,0.19]	0.145 [−0.02,0.23]	0.194 [0.03,0.26]
r_K^{UK}	0.130 [−0.09,0.26]	0.105 [−0.12,0.23]	0.101 [−0.10,0.27]	0.149 [−0.06,0.29]
r_{FF}^{USA}	0.000 [−0.14,0.11]	−0.088 [−0.20,0.04]	0.031 [−0.10,0.15]	0.041 [−0.10,0.14]
r_{DJ}^{USA}	−0.017 [−0.18,0.12]	−0.147 [−0.27,0.01]	0.035 [−0.07,0.21]	0.006 [−0.10,0.18]
r_{SP}^{USA}	0.026 [−0.11,0.12]	−0.076 [−0.19,0.05]	0.075 [−0.05,0.17]	0.070 [−0.06,0.16]

2) Augment 1960–2003 from 1970 onwards

	r_L^G	r_L^J	r_L^{UK}	r_L^{USA}
r_K^G	0.033 [−0.14,0.20]	0.133 [−0.05,0.24]	0.097 [−0.06,0.24]	0.154 [−0.01,0.26]
r_K^J	0.150 [−0.05,0.32]	0.167 [−0.07,0.28]	0.060 [−0.12,0.25]	0.178 [−0.02,0.32]
r_K^{UK}	−0.101 [−0.22,0.10]	−0.016 [−0.15,0.16]	0.138 [−0.07,0.22]	0.169 [−0.04,0.23]
r_{FF}^{USA}	0.021 [−0.13,0.15]	−0.00 [−0.16,0.10]	−0.069 [−0.18,0.09]	−0.033 [−0.16,0.11]
r_{DJ}^{USA}	0.069 [−0.16,0.18]	0.014 [−0.21,0.13]	−0.045 [−0.18,0.15]	−0.017 [−0.17,0.16]
r_{SP}^{USA}	0.057 [−0.12,0.15]	0.027 [−0.15,0.12]	−0.046 [−0.16,0.11]	−0.015 [−0.14,0.12]

3) Augment 1970–2012 from 2003 backwards

	r_L^G	r_L^J	r_L^{UK}	r_L^{USA}
r_K^G	0.118 [−0.08,0.23]	0.138 [−0.08,0.24]	0.110 [−0.07,0.26]	0.145 [−0.04,0.26]
r_K^J	0.0123 [−0.05,0.21]	0.071 [−0.09,0.19]	0.145 [−0.02,0.23]	0.195 [0.03,0.26]
r_K^{UK}	0.128 [−0.09,0.26]	0.105 [−0.12,0.23]	0.099 [−0.10,0.27]	0.147 [−0.06,0.29]
r_{FF}^{USA}	−0.001 [−0.14,0.11]	−0.088 [−0.20,0.04]	0.029 [−0.10,0.15]	0.039 [−0.10,0.14]
r_{DJ}^{USA}	−0.019 [−0.18,0.12]	−0.147 [−0.27,0.01]	0.087 [−0.07,0.21]	0.058 [−0.10,0.18]
r_{SP}^{USA}	0.024 [−0.11,0.12]	−0.076 [−0.19,0.05]	0.073 [−0.06,0.17]	0.068 [−0.07,0.16]

4) Augment 1970–2012 from 1970 backwards

	r_L^G	r_L^J	r_L^{UK}	r_L^{USA}
r_K^G	0.033 [−0.14,0.20]	0.133 [−0.05,0.32]	0.097 [−0.22,0.10]	0.154 [−0.14,0.13]
r_K^J	0.150 [−0.06,0.24]	0.168 [−0.07,0.28]	0.060 [−0.12,0.25]	0.178 [−0.02,0.32]
r_K^{UK}	−0.010 [−0.22,0.10]	−0.016 [−0.15,0.16]	0.137 [−0.07,0.22]	0.169 [−0.04,0.23]
r_{FF}^{USA}	0.022 [−0.14,13]	−0.000 [−0.16,0.10]	−0.069 [−0.19,0.08]	−0.033 [−0.17,0.10]
r_{DJ}^{USA}	0.070 [−0.16,0.18]	0.014 [−0.21,0.13]	−0.045 [−0.18,0.15]	−0.017 [−0.17,0.16]
r_{SP}^{USA}	0.058 [−0.12,0.15]	0.027 [−0.15,0.12]	−0.047 [−0.16,0.11]	−0.016 [−0.14,0.12]

Note: Correlations of human capital returns with physical capital returns. Returns to human capital are estimated using a VAR specification in levels with two lags. Returns on physical capital are computed using stock market data (CDAX for Germany, Nikko Securities Composite for Japan, FTSE All-Share Index for the United Kingdom, and Fama and French (FF), S&P 500 Total Return Index (SP), and Dow Jones Industrials Total Return Index (DJ) for the United States). We report in brackets the 95% confidence interval constructed using sampling-with-replacement raw residuals bootstrap based on 10,000 replications.

comparing the correlations vertically across tables. From a horizontal comparison across tables, however, one can see that the choice of the year in which we splice the datasets matters to some extent. Nevertheless, the main result of the table, remain unchanged: U.S. labor income returns innovations are always more correlated with foreign, rather than domestic, stock market returns (this is the key element for calibrating our buffer stock saving model).

Table A5
Value-weighted diversified portfolio with complete markets.

	Shares in each country traded asset:			
Investor nationality:	Germany	Japan	U.K.	U.S.
Panel A: r_k measured from VAR and national accounts				
Germany	−0.225 [−0.76,0.55]	0.060 −0.28,0.31]	1.576 [0.64,2.16]	−0.411 [−1.21,0.62]
Japan	0.742 [−0.19,1.76]	−0.606 [−1.08,−0.21]	1.882 [0.63,2.84]	−1.017 [−2.22,0.51]
UK	0.230 [−0.56,1.17]	1.059 [0.53,1.34]	1.191 [−0.08,2.02]	−1.480 [−2.50,0.08]
USA	0.019 [−0.52,0.91]	0.054 [0.11,0.78]	0.786 [−0.30,1.41]	−0.341 [−1.23,0.87]
Panel B: r_k measured using stock market returns				
Germany	0.040 [0.04,0.05]	0.289 [0.28,0.29]	0.145 [0.14,0.16]	0.526 [0.51,0.53]
Japan	0.034 [0.03,0.04]	0.290 [0.28,0.29]	0.139 [0.13,0.15]	0.537 [0.52,0.54]
UK	0.039 [0.03,0.05]	0.287 [0.28,0.29]	0.147 [0.14,0.17]	0.527 [0.5,0.53]
USA	0.039 [0.03,0.05]	0.287 [0.28,0.29]	0.145 [0.14,0.17]	0.530 [0.51,0.54]
World share	0.043	0.293	0.150	0.516

Note: Diversified portfolio using stock market returns (CDAX for Germany, Nikko Securities Composite for Japan, FTSE All-Share Index for the United Kingdom, and Fama and French for the United States). Each cell displays the net demand by a resident of country j for the assets of country k expressed as a fraction of home country (country j) marketable assets. We report in brackets the 95% confidence interval constructed using sampling-with-replacement raw residuals bootstrap based on 10,000 replications.

A.6. Hedging human capital in a frictionless complete market

The correlation structure between physical and human capital returns reported in Table 7 is not sufficient to identify the hedging behaviour without additional assumptions on the structure of the economy. In the main body of the paper, we provide a model of human capital hedging that accounts for both idiosyncratic and aggregate labor income risk, as well as liquidity constraints and market incompleteness. But in this subsection, we want to show that the VAR misspecification is the main driver behind the empirical results on human capital hedging reported in the previous literature. Moreover, we want to show that the estimates are characterized by substantial uncertainty, making the conclusion of human capital pushing toward overinvestment in foreign assets very far fetched. To show this, we now assume—only for this sub-section—that the set of international risky financial assets provides perfect spanning for human capital as in Bottazzi et al., 1996; Baxter and Jermann 1997. In other words, we assume that there exists a linear combination of domestic and foreign marketable assets that is perfectly correlated with the return to domestic human capital returns.

Assuming complete spanning of human capital returns by the stock market returns (as in Bottazzi et al., 1996; Baxter and Jermann, 1997), one can compute hedge portfolios for human capital risk via simple linear projection. Each hedge portfolio is chosen such that it hedges \$ 1.00 of human capital income flow. Let h_{jk} be the weight of the risky financial asset of country k in the hedge portfolio of country j residents. The hedge portfolio of country j, $h_j := [h_{j1}, h_{j2}, \ldots, h_{jK}]'$, is given given by

$$h_j = \Sigma^{-1}\Omega_j \tag{A1}$$

where Σ is the $K \times K$ covariance matrix of returns on risky financial assets in the world portfolio, and Ω_j is the $K \times 1$ vector of covariances of financial assets returns with the human capital return of country j. Notice that since the hedge portfolio is constructed to hedge \$ 1.00

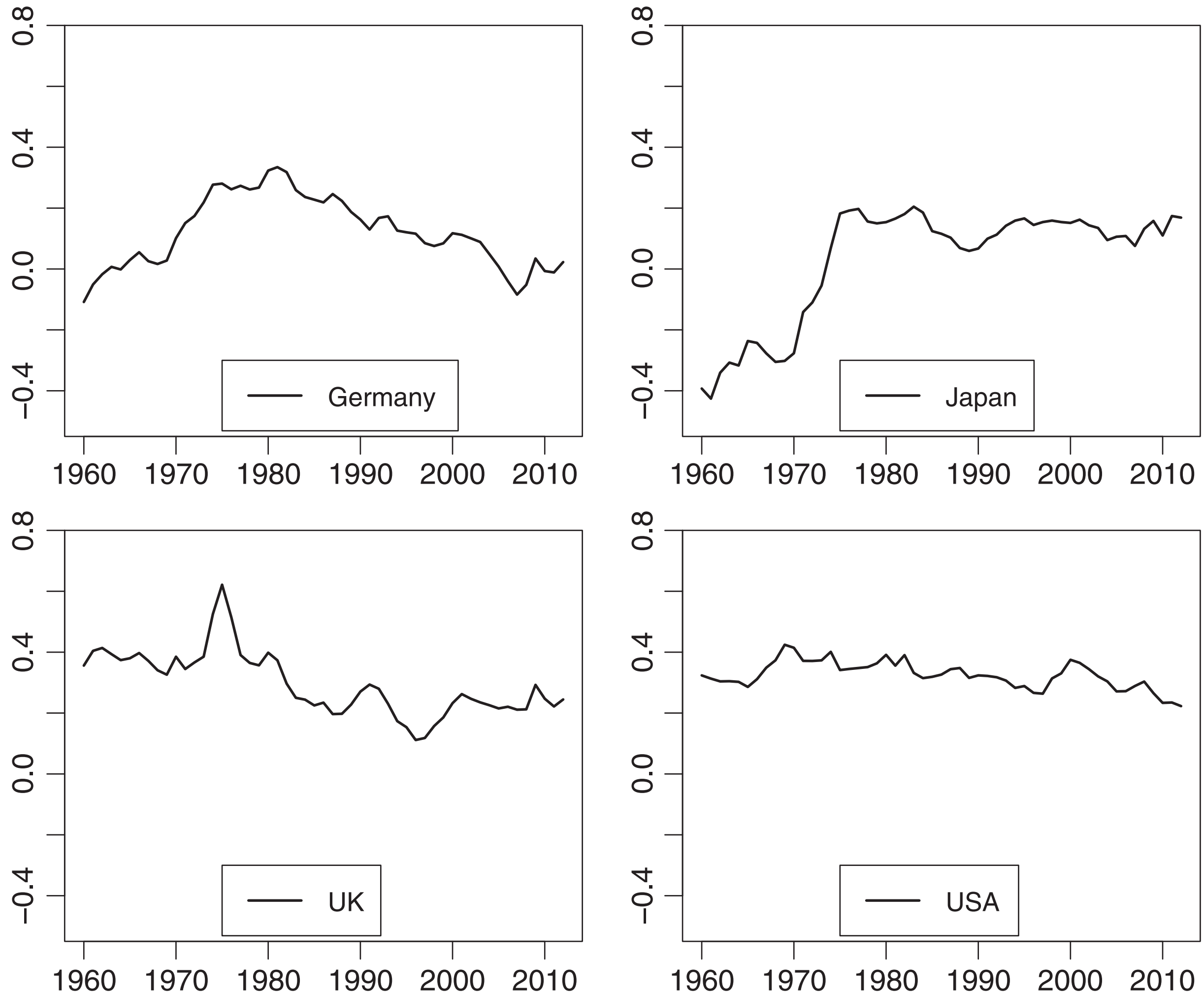

Fig. A4. Log ratio of labor income to capital income in Germany, Japan, the United Kingdom, and the United States.

of human capital, there is no reason for the portfolio weights to add to one.

Moreover, under the complete spanning assumption, one can also compute optimal portfolios following the value-weighted approach used in Baxter and Jermann (1997). The value-weighted portfolio approach, without considering the human capital hedging for the moment, follows from the two funds separation theorem (Merton, 1973): in a frictionless one-period economy with no asymmetric information, in the presence of K financial risky assets and a risk-free asset, all individuals will hold a portfolio given by a linear combination of the risk free asset and the market portfolio—i.e., the value-weighted portfolio. As a consequence, the risky part of each individual's portfolio will have a composition identical to the market portfolio. The extension of this approach to our international framework is straightforward: in equilibrium each investor, independent of his nationality, will hold a risky portfolio with a composition identical to the world portfolio of risky financial assets, i.e., each country's asset will be in the portfolio with a share equal to the share of the country in the world portfolio of marketable risky assets. That is, denoting with π_k the fraction of country k risky marketable asset in the world portfolio, we have that in the absence of human capital risk, π_k would be the share of country k asset in the portfolio of risky financial assets of each country.

In the presence of non-marketable human capital, portfolio holdings will be adjusted due to the hedging motive. Therefore, the portfolio of risky assets held by a country j individual will depend also on the covariance of domestic returns to human capital and (domestic and foreign) returns to physical capital. In particular, to hedge the human capital risk, the net demand by a resident of country j for the asset of country k expressed as a fraction of the home country (country j) portfolio of financial assets will be

$$\pi_k\left[1+\frac{1-\alpha_j}{\alpha_j}\left(\sum_{k=1}^{4}h_{jk}\right)\right]-\frac{1-\alpha_j}{\alpha_j}h_{jk}, \tag{A2}$$

where $1-\alpha_j$ is the labor share of income in country j.

The last term in Eq. (A2) is the share of country k marketable asset that has to be sold to hedge the human capital risk in country j. The expression multiplying π_k reflects the funds generated by selling the investor's endowment of the claim to domestic physical capital and the portfolio that hedges the risks associated with human capital wealth, i.e., $\frac{1-\alpha_j}{\alpha_j}\left(\sum_{k=1}^{4}h_{jk}\right)$.

The value weighted optimal portfolios implied by the correlations between human and physical capital returns in Table 8 are reported in Table A5. Panel A focuses on the VAR and national accounts based estimates of physical capital returns, while panel B use stock market index returns.

Panel A shows that, after correcting the VAR misspecification, the Baxter and Jermann (1997) value-weighted approach leads to inconclusive results concerning the role of human capital for optimal portfolio diversification: large home bias is basically as likely as large shorting of the domestic market.

Comparing the optimal in portfolios in panel B with the (rescaled to sum up to one) world shares (π_k) of the various countries' stock markets in the world market capitalization reported in the last row, the table shows that the inclusion of human capital hedging motive has basically no effect when the space of tradable claims to capital is the one of publicly traded companies.

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