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 show that, given the high degree of international GDP correlations in the data, very small domestic redistributive shocks can skew portfolios toward domestic assets. Second, we find that the empirical evidence supporting the belief that the international diversification puzzle is worsened if we consider the human capital hedging motive, is driven by an econometric misspecification rejected by the data. Third, we find that if the set of assets that can be used to hedge aggregate human capital is restricted to publicly traded stocks, the human capital hedging motive has a negligible impact on optimal portfolio choice in a frictionless market setting. Fourth, calibrating an heterogeneous agent model in which households face short selling constraints and labor income risk, in the form of both uninsurable shocks and a common aggregate component, we show that: *a)* investors that enter the stock market will initially specialize in domestic assets, *b)* individual portfolios become gradually more internationally diversified only as the level of asset wealth increases, and *c)* most importantly, the implied aggregate portfolio of domestic investors shows a large degree of home bias.

Keywords: international diversification puzzle, nontraded human capital, hedging the returns to human and physical capital, optimal portfolio choice
JEL classification: F30, G11, G12.

"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).

I Introduction

International finance theory emphasizes the effectiveness of global portfolio diversification strategies for cash-flow stabilization and consumption risk sharing.[1] However, the empirical evidence on international portfolio holdings concludes in favour of a widespread lack of diversification across countries, and a systematic bias toward home country assets. This discrepancy between theoretical predictions and observed portfolio constitutes the *international diversification puzzle* (see, e.g., Fama and French (1991) and Sercu and Vanpe (2007)).

In the major industrialized countries, roughly two-thirds of Gross Domestic Product goes to labor and only one-third to capital. Thus human wealth is likely to be about two-thirds of total wealth. This suggests 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.

In this paper we ask whether the human capital hedging motive is likely to have a sizeable effect on optimal portfolio choice, and what are its implications for the international diversification puzzle.

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.[2] 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, Pesenti, and van Wincoop (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

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[1]Nevertheless, the size of gains from international risk sharing continues to be a debated issue. 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. Obsfteld (1994) calibration exercises imply that most countries reap large steady-state welfare gains from global financial integration.

[2]Moreover, Michaelides (2003) shows with a calibration exercises 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.

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.

This paper offers two main sets of contributions. First, we show that the canonical intuition that human capital should skew portfolio holdings toward foreign assets, and the related supporting empirical evidence, are both very fragile. Once the degree of international economic integration observed in the data is properly taken into account, extremely small redistributive shocks can lead to a home country bias in portfolio holdings. This result arises even in a setting in which, in the absence of small redistributive shocks, payoffs to domestic human and physical capital would be perfectly correlated. Moreover, 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.

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 VAR approach. This feature of the data provides a rationalization for the apparently contradicting empirical evidences on the correlation between returns to human and physical capital found in the previous literature: Lustig and Nieuwerburgh (2008) finds a strong negative correlation between returns to physical and human capital in U.S. data; Bottazzi, Pesenti, and van Wincoop (1996) finds such a correlation to be negative in all the country they consider but the U.S.; Baxter and Jermann (1997) find this correlation to be positive and very close to one for all the countries they consider (U.S., U.K., Germany and Japan). Nevertheless, we find that 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. 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 that the effect is quantitatively small in a frictionless complete market setting.[3]

Second, using the estimated correlations between human capital and stock market

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

returns innovations uncovered in the data, we calibrate an incomplete market model in which agents face both idiosyncratic and aggregate labor income risk, as well as borrowing constraints. 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 low levels of liquid wealth will initially specialize in domestic assets and, b) only at large ratios of asset wealth to labor income agents significantly diversify their portfolios internationally, c) as a consequence, the aggregate portfolio of U.S. investors shows a large degree of home bias.

What drives this result? First, households face large human capital risk, but this is mostly of the idyosincratic type – hence underestimated in an homogeneous agent setting. Second, 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 lowest 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 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, overall, our results 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&P500 is substantially negative for some categories.[4] Moreover, they find that for six out of the eight sex-education groups

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[4] Davis and Willen (2000) generally finds that the degree of correlation between earning shocks and equity returns rises with education, with a lower bound correlation of -.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 complimentary to physical capital and the use of advanced technologies.

considered in their study, a *long* position on the worker's own industry represent a good hedge for labor income risk. The empirical works of Gali (1999), Rotemberg (2003) and Francis and Ramsey (2004) also document a negative correlation between labor hours and productivity conditioning on productivity shocks.

The analysis presented in this paper is part of, and complementary to, the literature that has attempted to explain home bias as a hedge against non-tradable risks.[5] Moreover, the potential rationalization of the international diversification puzzle we domucent in this paper should be interpreted as complementary to the ones based on information frictions (Nieuwerburgh and Veldkamp (2009)), nominal stickiness (Engel and Matsumoto (2009)), and the role of non-traded goods (Heathcote and Perri (2013)).

The remainder of the paper is organized as follows. Section II shows how very small rent shifting shocks can lead to home country bias in portfolio holdings where the degree of international correlations in GDP movements is taken into account. Section III presents the approach undertaken to measure factor returns. Empirical results are discussed in Section IV while Section 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. Conclusions are outlined in the final section.

II   Rent Shifting Shocks and the Home Country Bias

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.

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.[6]

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 $\delta = corr\left(y, y^f\right)$ the correlation between domestic and foreign output (y^f). It then follows that

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

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

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

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

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 $\alpha = .3$, Figure 1 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

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[6]Many kinds of shocks are expected to have an effect on the dynamics of income distribution. 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 U.S. 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 U.S., finds a large and negative correlation between unexpected union wage changes and unexpected changes in the stock value of the firm; Bottazzi, Pesenti, and van Wincoop (1996) provide evidence of a negative correlation between wage and profits rates in many OECD countries; Lustig and Nieuwerburgh (2008) uncover a negative correlation between innovations to human and physical capital returns in the U.S.; Gali (1999), Rotemberg (2003) and Francis and Ramsey (2009) document a negative correlation between labor hours and productivity conditioning on productivity shock.
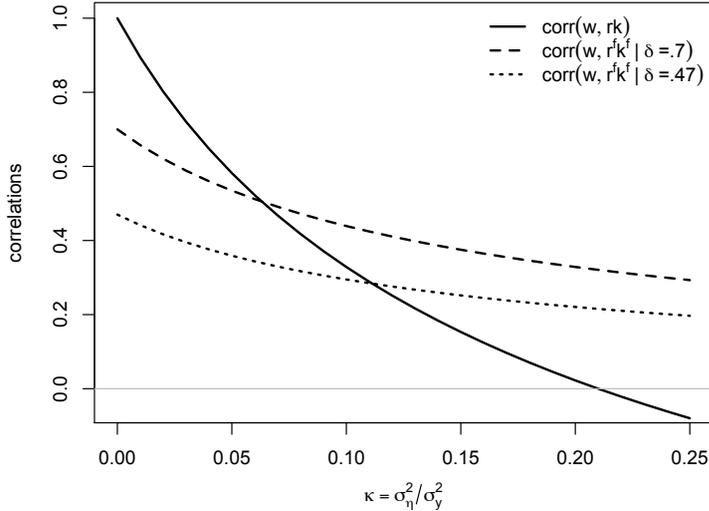
Figure 1: Correlations of domestic wage and: domestic (solid line) and foreign (dashed and dotted lines) capital compensation for different levels of international GDP correlation (δ).

variance to output variance). The two values of δ considered are the minimum and maximum correlations between domestic output and world output reported in Obsfteld and Rogoff (1996, page 291) for Germany, UK, US and Japan.[7] 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 .7 (.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 (US, UK, Japan and Germany during the 1960–2012 period)[8] that ranges from .06 to .48, when using Hodrick and Prescott (1997) de-trending, and from .09 to .18 when using linear de-trending. That is, these estimate are in the range needed to skew portfolio holdings toward domestic assets.[9] Moreover, in the presence of rent shifting shocks, we would

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

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

[9]We find very similar results if we gauge the magnitude of κ by computing the ratio of i) the

expect to observe time variation in the labor share of income. Figure 2 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.
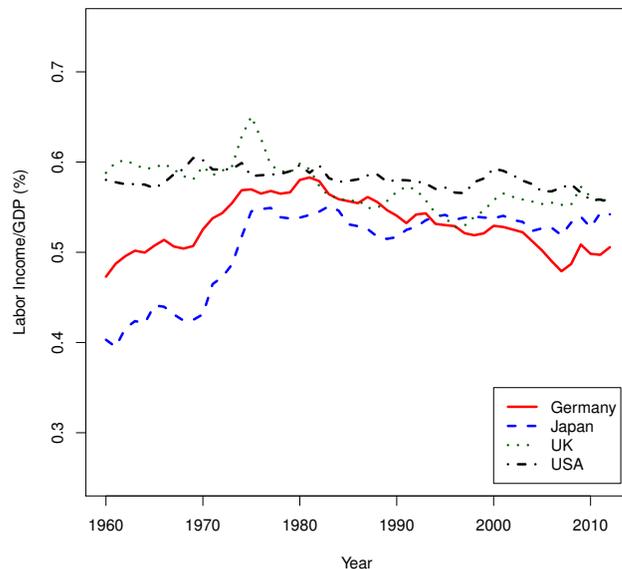
Figure 2: Labor share of income in Germany, Japan, UK and USA

III Measuring Factor Returns

To assess the role of human capital in determining optimal portfolio choice, we need 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

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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.

asset. The gross (unobserved) return R is given by the following accounting identity:

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

Assuming that the price-dividend ratio is stationary, we can log-linearize Equation (1)) 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

$$p_t - d_t = \sum_{\tau=1}^{\infty} \rho^{\tau-1} \left[\Delta d_{t+\tau} - r_{t+\tau} \right] + \lim_{T \to \infty} \rho^T \left(p_{t+T} - d_{t+T} \right). \tag{2}$$

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 Equation (2) holds for any realization of $\left\{ \Delta l_{t+\tau}^e - r_{h,t+\tau}^e \right\}_{\tau=1}^{\infty}$ and $p_\infty - l_\infty$, and hence holds in expectation for any probability measure. This implies that we can take expectations of Equation (2) under both the 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,[10] i.e. $\lim_{T \to \infty} \rho^T \left(p_{t+T} - d_{t+T} \right) = 0$, we have that

$$r_{t+1} - \bar{r} = \sum_{\tau=1}^{\infty} \rho^{\tau-1} \left(E_{t+1} - E_t \right) \left[\Delta d_{t+\tau} \right] \tag{3}$$

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. Equation (3) 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 Equation (3). We perform this task following Campbell and

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[10]Imposing that the transversality condition holds is less restrictive than it might seem since, even thought 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)).

Shiller (1988) and use linear projections generated by a reduced form vector autoregressive model (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) we focus on four coutries – U.S., U.K., Germany and Japan[11] – and the variables included in our VAR are the labor and capital incomes in each of this countries. But, differently from Baxter and Jermann, our empirical procedure allows for common international shocks, comovements, and trends among countries and, as show in Section IV, this difference will lead 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 vector autoregressive 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. Their approach embeds the assumption of low international economic integration. Their VECM, for each country i, takes the form:

$$
\begin{bmatrix} \Delta d^i_{L,t+1} \\ \Delta d^i_{K,t+1} \end{bmatrix} = \begin{bmatrix} c^i_L \\ c^i_K \end{bmatrix} + \begin{bmatrix} \psi^i_{LL}(L) & \psi^i_{LK}(L) \\ \psi^i_{KL}(L) & \psi^i_{KK}(L) \end{bmatrix} \begin{bmatrix} \Delta d^i_{L,t} \\ \Delta d^i_{K,t} \end{bmatrix} + \begin{bmatrix} \eta^i_L \\ \eta^i_K \end{bmatrix} \left(d^i_{L,t} - d^i_{K,t} \right) + \begin{bmatrix} \varepsilon^i_L \\ \varepsilon^i_K \end{bmatrix}
$$
(4)

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 rational 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 will reach 1 or 0 with probability 1 as the sample size goes to infinity. Section A.2 of the Appendix reports a formal empirical analysis of this assumption.

Equation (4) can be rewritten in more compact form as:

$$
\Delta D^i_{t+1} = C^i + \Psi^i(L)\Delta D^i_t + \Pi^i \left(d^i_{L,t} - d^i_{K,t} \right) + \nu^i_{t+1}
$$
(5)

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

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[11]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%.

for each of the four countries considered, the VECM estimated by Baxter and Jermann (1997) can be rewritten as a system of the form

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

(6)

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

There are two important implicit restrictions in Equation (6). 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 supposed *not* 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 Equation (6) 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 section relaxes both of these restrictions, and considers a more general class of vector autoregressive models for labor and capital incomes, that allow for short- and long-run comovements across countries.

IV Empirical Evidence: A Reappraisal

Equations (3) and (6), as well as alternative VAR specifications, are estimated using annual data on labor income and capital income from OECD National Accounts for Germany, Japan, United Kingdom and United States over the period 1960-2012. Our measure of labor income is total employee compensation; 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.

IV.1 Model Selection

The restrictions imposed in Equation (6) 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 1 reports frequentist tests of the null hypothesis of block exogeneity in Equation (6). 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 Equation (6). 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 suggests that there exist statistically significant cross-country links between the compensations of the factors of production.

Table 1: 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 Equation (6). p-value of not rejecting the null hypothesis in brackets.

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.[12] As a consequence, we perform model comparison employing Bayesian posterior probabilities (see e.g. Gelman, Carlin, Stern, and Rubin (1995)) since this approach is robust to non-stationarity in the data. In particular, for each model j considered we compute the Bayes factor

$$BF_j = \int_{\Theta_j} g_j(\theta) p_j(X \mid \theta) d\theta \tag{7}$$

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 j-th model. Using the Laplace method (see e.g. Schervish (1995)) the Bayes factor can be approximated as

$$BF_j \cong g_j(\widehat{\theta}_j) p(X \mid \widehat{\theta}_j)(2\pi)^{\frac{m}{2}} \left| \hat{\Sigma}_{\theta j} \right|^{\frac{1}{2}} \tag{8}$$

where $p_j(X \mid \widehat{\theta}_j)$ is the likelihood of the j model evaluated at its peak $\widehat{\theta}_j$, 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 be

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

computed as

$$PO_j = \frac{p_j BF_j}{\sum_i p_i BF_i} \tag{9}$$

where p_j is the prior probability of the j-th model.

Table 2 reports the logs of the Bayes factor and the posterior probabilities defined by Equations (8) and (9) for a large set of models, under the assumption of flat priors and equal prior probability for each model. The models considered are: 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 cointegration 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; iv) unrestricted VARs in levels (rows 7 and 8) which 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.

Table 2: 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.

The econometric model considered in row 1 of Table 2 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 fac-

tor ($\log BF_i$). 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 – 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 1. 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 than the one in the the first row, but much smaller than the one in the 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 provides supporting evidence for within country cointegration in the payoffs to physical and human capital. Nevertheless, in Section A.2 of the Appendix 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 the rows 7 and 8 of table 2 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 2) that is numerically undistinguishable 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 Equation (7) (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 2 and we also experimented with non-flat priors over the parameters space. In both cases, the results are in line with the ones in Table 2.

Overall, the results of this section imply that to accurately measure returns to human and physical capital, and their implications for international portfolio diversifica-

tion, 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. As a consequence, in the next subsections we construct hedging portfolios and optimal portfolio holdings using such an econometric specification.

IV.2 The Correlation of Human and Physical Capital Returns

In order to compute estimate factor returns using Equation (3) 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 Equation (3) 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 3 reports the correlations between returns on capital and labor computed using Equation (3) and the estimations of expected Δd's by the VAR in levels specification with two lags.

Table 3: 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.

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 [0.78, 0.99], our estimates have a maximum of 0.96 and a minimum of 0.67 in Japan.[13]

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[13]Bottazzi, Pesenti and van Wincoop (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).

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 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.).

Instead 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 international level.

The factor returns that we obtain by applying Equation (3) 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 for instance 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 3, 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 Baxter and Jermann (1997) results are mostly driven by the relaxation of the block exogeneity assumption, which turns out to be strongly rejected by the data. They 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.

For comparability, in Table A5 of Appendix A.5 we replicate the optimal portfolio implied by the complete spanning approach of Baxter and Jermann (1997). That is, we replicate the author's main result, but correcting the VAR misspecification. Raw point estimates in the table implicates that the authors original claim, that the home country bias is generally worsened by the human capital hedging motive since, the table shows, for some countries we obtain the opposite effect as Bottazzi, Pesenti and van Wincoop (1996). On the other hand, confidence intervals show that there is substantial uncertainty about optimal portfolio, and that human capital hedging can potentially generate large home country bias. This finding also reconciles the discrpancy between

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the optimal portfolios estimated in Baxter and Jermann (1997) and Bottazzi, Pesenti and van Wincoop (1996).

IV.3 The Correlation of Human Capital and Stock Market Returns

Estimating returns to physical capital using the approach in Equation (3) 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 (relevant for hedging purposes) correlation between human capital compensation and physical capital ones, 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 sections.

Table 4 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 will play a key role in the model calibration in the next section, for the United States we use three different stock indices: the Fama and French (1992) benchmark market return,[14] 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 U.S. 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.4, 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

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[14]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/`

Table 4: 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 United States). We report in brackets the 95% confidence interval constructed using sampling-with-replacement raw residuals bootstrap based on 10,000 replications.

of optimal portfolios as in Baxter and Jermann (1997), the effect of human capital hedging would be very small – but would tend to skew holdings in favour of domestic assets, as shown in Table A5 in Appendix A.5.

Nevertheless, as shown in the next section, 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.

V Portfolio Choice with Heterogeneous Human Capital Risk and Liquidity Constraints

In this section we relax the simplifying assumptions needed in the previous section to obtain close form solutions for hedge portfolios and portfolio holdings, and 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 model presented below is a generalization of Heaton and Lucas (1997) to a multiple asset context and of Michaelides (2003).

V.1 Model Set-up 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

$$C_t + B_t + S_t^d + \sum_{j=1}^{J} S_t^j \leq R_t^f B_{t-1} + R_t^d S_{t-1}^d + \sum_{j=1}^{J} R_t^j S_{t-1}^j + Y_t, \tag{10}$$

and the standard transversality condition, where $1 > \beta > 0$ is the discount factor (calibrated at the value of .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 is 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.

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.3 for details) an MA(2) specification for its log growth rate:

$$g_{t+1} = \log \frac{Y_{t+1}^g}{Y_t^g} = \mu_y + \varepsilon_{t+1} + \theta_1 \varepsilon_t + \theta_2 \varepsilon_{t-1} \tag{11}$$

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

$$Y_t^i = Y_t^g P_t^i U_t^i \tag{12}$$

$$P_t^i = G P_{t-1}^i N_t^i \tag{13}$$

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 $E\left[U_t^i\right] = 1$, $\log P_t^i$ 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 correspond to Gourinchas and Parker (2002) except for the added term Y_t^g that reflects aggregate economic uncertainty.[15] Following Gourinchas and Parker (2002) estimates, we calibrate $\sigma_u = 0.073$ and $\sigma_n = 0.105$, and following we calibrate σ_ε, θ_1, θ_2 and μ_y using the point estimates in Table A3 in Appendix A.3. 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 equations (11)-(13) the individual labor income growth is given by

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

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

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 in equation (13) 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 satiations in consumption, the budget constraint will hold with equality and

$$C_t = X_t - 1_{\{B_t > 0\}} B_t - 1_{\{S_t^d > 0\}} S_t^d - \sum_{j=1}^{J} 1_{\{S_t^j > 0\}} S_t^i \tag{14}$$

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

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

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 equation (14) and the homogeneity of degree $-\gamma$ of the marginal utility, we can rewrite the Euler equations as

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

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

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

$$\forall j' = 1, \dots J$$

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

$$x_t = \left(R^f b_{t-1} + R_t^d s_{t-1}^d + \sum_{j=1}^{J} R_t^j s_{t-1}^j\right)\frac{Z_{t-1}^i}{Z_t^i} + \frac{Y_t^i}{Z_t^i} \tag{15}$$

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

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

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

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

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

As in the previous sections, and to avoid the curse of dimensionality, we only focus on four countries: the U.S. (as domestic country), U.K., Japan and Germany, since the market capitalization of this four countries represents close to ninety percent of the world market.

Since the U.S. domestic risky asset has enjoyed both the lowest variance and a high 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 U.S.. A summary of the calibrated preference and labor income process parameters reported in Table 5.

Table 5: Preference and Labor Income Parameters

γ	3
β	0.95
σ_U	0.105
σ_N	0.073
μ_y	0.019
ϑ_1	0.448
ϑ_2	0.094
Mean Market Return	0.060
Market Return sd	0.175
Risk Free Rate	0.011

The crucial element in calibrating the model is the covariance structure of asset returns and innovations to the aggregate labor income process. As in Table 4 we measure capital returns using broad stock market indexes, and calibrate their covariance using the time series sample analogous. To calibrate the covariance between aggregate labor income shocks and stock market returns, we use the estimates reported in Table 4. These correlations are all small in magnitude and, as shown in Table A5, have a small effect on the optimal portfolio choice in the absence of short-selling constraints.

Table 6: Market Returns and Aggregate Labor Income Shock Correlations

	Correlations				Implied market
	Germany	Japan	U.K.	Aggregate labor income shocks	portfolio w.o. labor income risk
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 calibration of the covariance structure is summarized by the correlations reported in the first four columns of Table 6. As a benchmark, the last column of the table also reports the implied optimal portfolio shares in the market portfolio absent any human capital hedging motive and shows that, according to the estimated market return correlations, the share of U.S. assets in the U.S. domestic portfolio should only be of the order of 25 percent in the absence of aggregate labor income risk.

V.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) and Haliassos and Michaelides (2002)) to compute the optimal consumption and asset holding rules.

Since 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,[16] obtaining the investment rules $\{b(x), s^d(x), s^i(x)\}$. Moreover, from equation (14) we can obtain the optimal consumption rule $c(x)$.

Optimal policy rules are plotted, as a function of normalized cash-on-hand, in Figure 3. 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 know 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.
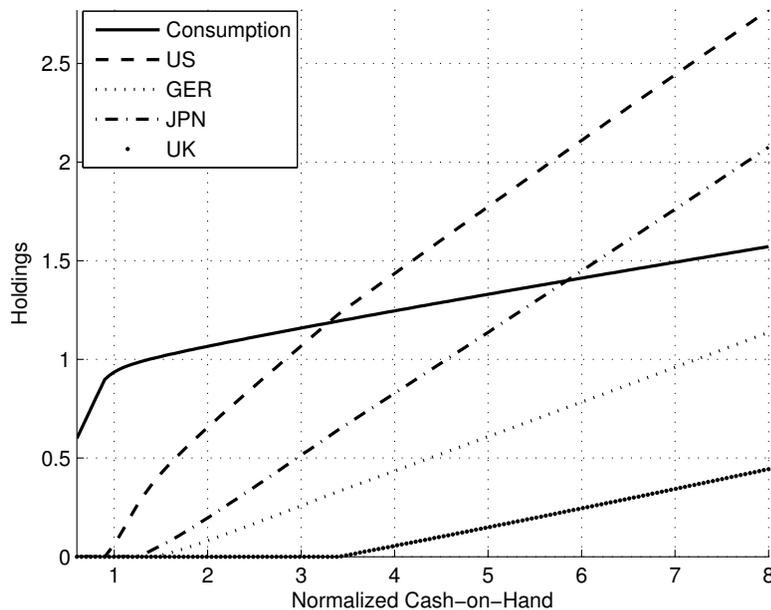

Figure 3: Optimal consumption and investment policy functions as a function of normalized cash-on-hand.

More interestingly, when the consumer enters the saving region she initially invests only in the domestic stocks and, only as the level of cash on hand increases, does she gradually diversify her portfolio internationally. 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

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[16]Optimal policy functions do not seem to change significantly as a function of past aggregate labor income shocks, manly due to the very small variance of these shocks compared to the idiosyncratic ones. In particular, policy function 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.

to invest in the assets that has 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 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 results 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 the four assets considered. Moreover, given the positive correlation between normalized cash-on-hand and 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, Julliard, and Pagano (2001)).
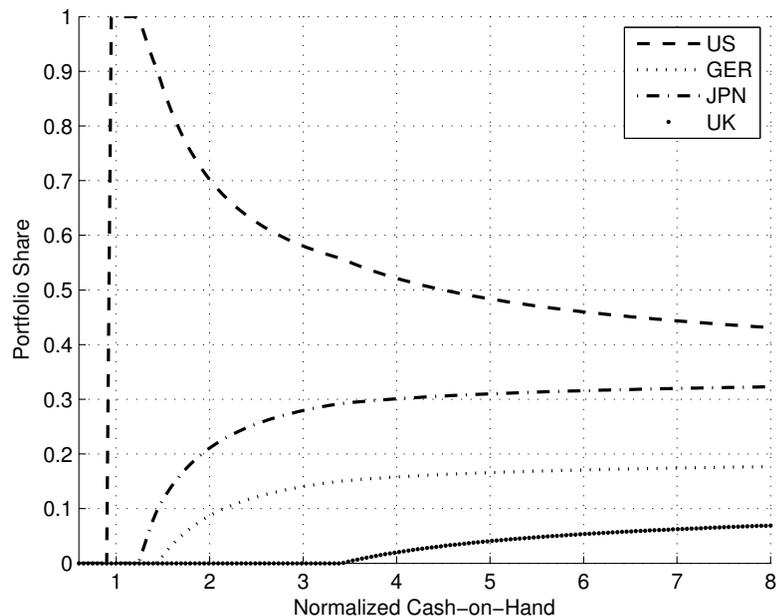

Figure 4: Optimal portfolio shares as a function of normalized cash-on-hand.

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 Figure 4. The

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figure shows a large bias toward domestic assets in all the relevant range of standardized cash on hand implying that more than 99 percent 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 percent in Table 6), this represent a home country bias of individual portfolios that ranges from 75 to 19 percent. Even investors in the top one percent of the distribution of cash-on-hand observed in the data, would have on average more than 50 percent of their asset wealth invested in domestic stocks. Interestingly, this large effect 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 the counterfactual calibration results in Appendix A.6, 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 labor income innovations and market returns at a 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.

V.3   Implications for the Aggregate Portfolio

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

The standardized cash-on-hand in equation (15) follows a renewal process and can be shown to have an associated invariant distribution,[17] 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 equation (15) to compute, by repeated simulation over a grid of values, the transition probabilities from one level of cash-on-hand to the other

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

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

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[17]See, e.g., Deaton and Laroque (1992), Carroll (1997), Szeidl (2002) and Carroll (2004).

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.
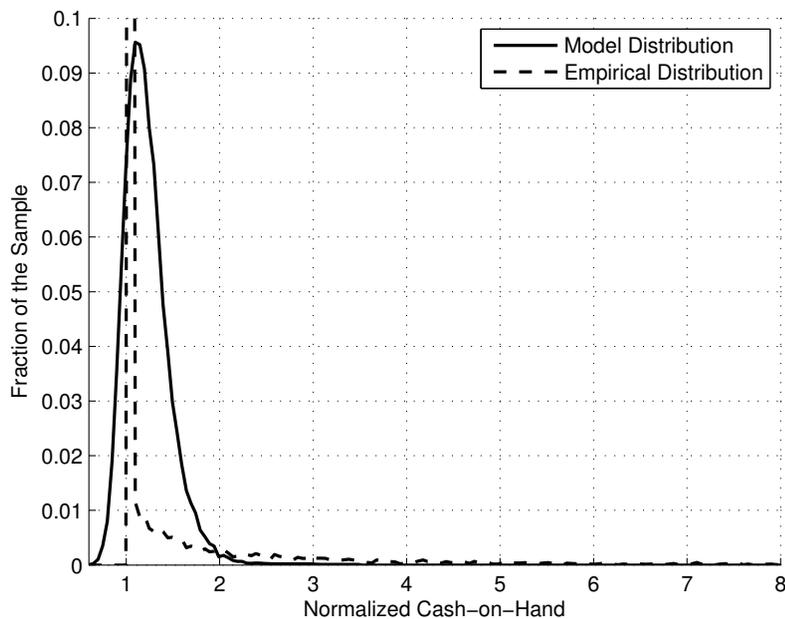

Figure 5: Model implied and empirical invariant distributions.

Figure 5 reports the distributions of normalized cash-on-hand implied by the model and the observed distribution of normalized cash-on-hand in the Panel Study of Income Dynamics ($PSID$) data.[18] 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

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[18]The $PSID$ data contains accurate information on wealth holding of household at five years 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.

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 set up 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 7 reports, as 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.

Table 7: Aggregate Portfolio Shares of U.S. Investors with Liquidity Constraints

	No Human 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%

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 7 could be a poor approximation of the aggregate portfolio. To address this issue, the third column of Table 4 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 percent, implying that hedging human capital increases the portfolio share of domestic stocks by as much as 50 percent and decreases the portfolio shares of German, Japanese and U.K. stocks by, respectively, 15, 19 and 17 percent.

The last two columns of Table 7 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 case, quite similar to the ones obtained 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 percent.

But what is the key mechanism delivering a large home bias generated by the model in Table 7? The driving force of our 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.

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.

This implies 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. Bottazzi, Pesenti, and van Wincoop (1996), using a VAR approach that once again imposes block exogeneity across countries (hence is likely to overestimate the benefits of international portfolio diversification), finds that the correlations of returns to human capital with domestic market returns is smaller than the one with foreign market returns in seven out of ten countries in their study (with an average difference of .19). Therefore, the human capital hedging motive is likely to explain a large fraction of the home country bias in several countries.

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.[19] 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.

VI   Conclusion

This paper shows that human capital risk does not seem to be a compelling reason to reduce the share of domestic assets in the portfolios of domestic investors. We provide four main results pointing in this direction.

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: very small redistributive shocks – e.g. with a variance of a mere 6% of GDP variance – are enough to reverse this intuition. Moreover, the presence of this type of shocks is strongly supported by the data.

Second, we show that the commonly used approach of estimating country specific Vector Autoregressions 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 mispecification largely drives the findings of Baxter and Jermann (1997), i.e. the result that the home bias puzzle is uneqiovocally worse than we think once we consider human capital risk, is the outcome of an econometric misspecification strongly rejected by the data. Our results show that the evidence in the data is much more nuanced: first, removing the mispecification, human capital hedging helps explain the home country bias in some countries (as in Bottazzi, Pesenti, and van Wincoop (1996)); second, we show that the approach used in the previous litterature leads to very large confidence intervals, rendering point estimates largely uninformative.

Third, we show that when measuring returns to capital using broad stock market indexes, return to human capital tend to be more correlated with foreign rather than domestic stocks, consistently with the rent shifting hypothesis. Nevertheless, these correlations are very small, hence in a frictionless complete market setting have very little impact on optimal portfolios.

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 can be obtained as an equilibrium result. This is due to the fact 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

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[19]Hau and Rey (2003) finds that, at higher frequencies, higher returns in the home equity market relative to the foreign equity market are associated with a home currency depreciation.

correlation with the labor income innovations – the domestic asset.

Our findings should not be interpreted as alternative to other rationalisation of the home country bias in portfolio holdings (see, e.g., Engel and Matsumoto (2009), Nieuwerburgh and Veldkamp (2009), Heathcote and Perri (2013)) but rather complementary.

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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, is 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 which 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. Therefore, 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 choose to augment the dataset from 1960 until 2003 with more recent data which is adjusted using the GDP ratio for the year 2003 as conversion rate. Section A.4 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 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 of income reported in the table are the values for α_j used in Equation (17).

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, S&P 500 Total Return Index and Dow Jones Industrials Total Return Index (DJ) for 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 and GDP figures deflator from the International Financial Statistics service of the International Monetary Fund (available through `http://ukdataservice.ac.uk/`).

Table A1: Summary Statistics

		Labor Income	Capital Income	Labor Share
Germany	Mean	190943.24	170546.87	0.530
(millions of euros)	Std. Deviation	62059.31	63631.03	0.029
Japan	Mean	154322.42	121684.88	0.508
(billions of yen)	Std. Deviation	62670.50	46027.72	0.047
U.K.	Mean	70718.70	53746.65	0.573
(millions of pounds)	Std. Deviation	19304.07	18159.68	0.025
U.S.	Mean	150265.45	109329.31	0.580
(millions of dollars)	Std. Deviation	43485.35	34474.46	0.011

Note: The sample period is 1960–2012.

A.2 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 A2 presents econometric evidence based on the Johansen cointegration test (Johansen (1995)), where 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 A2 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.

Table A2: 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.

In Figure 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 A2, formal testing procedure based on the Johansen cointegrating test strongly rejects the cointegration restriction for all countries.
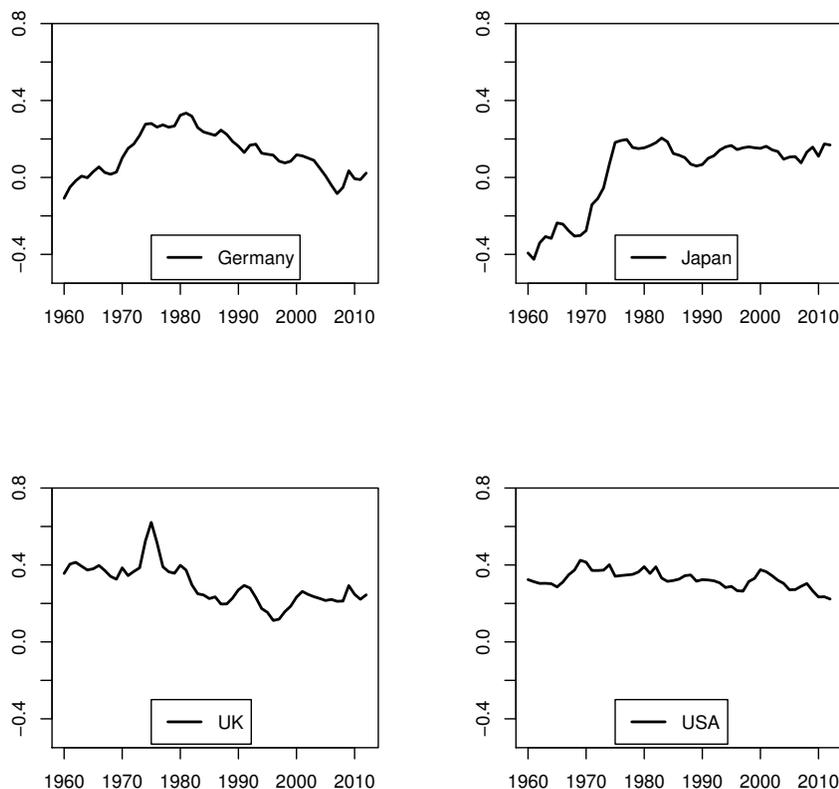

Figure A1: Log ratio of labor income to capital income in Germany, Japan, U.K. and the U.S.

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 it is completely silent about the existence of a long-run equilibrium relationship across countries. Moreover, since in the paper we do our econometric analysis in levels, we allow for implicit cointegrating relationships in the data.

A.3 Estimation of the Aggregate Labor Income Process

In order to model the U.S. labor income process, we experimented with several specification in the ARIMA class, and performed the standard set of Box-Jenkins selection procedures. In particular, among the model considered, MA(2) and ARMA(1,1) processes fit well to first differences of log labor income. 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 used in the literature in similar contexts (see, e.g., Davis and Willen (2000) and MaCurdy (1982)). Thus, the fitted earning specification is the MA(2) process in equation (11). The estimated coefficients are reported in Table A3 below.

Table A3: Estimated Labor Income Process

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

Note: Newey-West standard errors reported in brackets.

A.4 Robustness of Factor Returns Correlations

This section tests whether the correlation of the factor returns presented in table 4 are sensitive to the way we construct our data series on labor income. As mentioned in section A.1, our final data of labor and capital income stems from two different datasets which are then combined. An immediate concern is whether the results of the 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 datset from 1970 backwards. We opt for the first alternative for several reasons. First, we choose to augment the 1960-2003 dataset since this is 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 has already reunified and the data is thus consistent. Nevertheless, we explore the robustness of our results by comparing them for the four different alternatives.

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 United Kingdom, Fama and French (FF), S&P 500 Total Return Index (SP) and Dow Jones Industrials Total Return Index (DJ) for United States). We report in brackets the 95% confidence interval constructed using sampling-with-replacement raw residuals bootstrap based on 10,000 replications.

In table A4 we recalculate table 4 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 wether we extend the 1960-2003 or the 1970-2012 dataset. This can be seen from comparing the correlations vertically across tables. From a horizontal comparison accross 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 which is key for this paper remains unchanged: the ordering of the correlations of the different stock market returns with the U.S. labor returns. Even though the numbers change, it is still true that this correlation is largest for Japan, followed by the U.K. and Germany and, most importantly, it is lowest for the U.S.

A.5   Hedging Human Capital in a Frictionless Complete Market

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The correlation structure between physical and human capital returns reported in Table 3 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 subsection – that the set of international risky financial assets provides perfect spanning for human capital as in Bottazzi, Pesenti, and van Wincoop (1996) and 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 as in Bottazzi, Pesenti, and van Wincoop (1996) and 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, $h_j := [h_{j1}, h_{j2}, ..., h_{jK}]'$, is given given by

$$h_j = \Sigma^{-1}\Omega_j \tag{16}$$

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 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, independently 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

$$\pi_k \left[1 + \frac{1 - \alpha_j}{\alpha_j} \left(\sum_{k=1}^{4} h_{jk} \right) \right] - \frac{1 - \alpha_j}{\alpha_j} h_{jk}, \tag{17}$$

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

The last term in Equation (17) is the share of country k marketable asset that has to be sold to hedge the human capital risk in country j. The expressions 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 4 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.

Table A5: Value-weighted diversified portfolio with complete markets

Investor Nationality:	Shares in each country traded asset:			
	Germany	Japan	UK	USA
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 United Kingdom, and Fama and French for 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.

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.

A.6 A Counterfactual Calibration

This section shows that the results presented in section V.2 are almost entirely driven by the relative ordering of correlations between labor income innovations and market returns, rather than the particular point estimates.

To stress this point, we perform a counterfactual calibration of the model. First, to focus only on the effect of the correlations of labor income and asset returns, we

calibrate the between assets correlation to be all equal to .5. Second, we consider four assets that have equally spaced positive correlations with labor income innovations that range from 0.0 to 0.15.
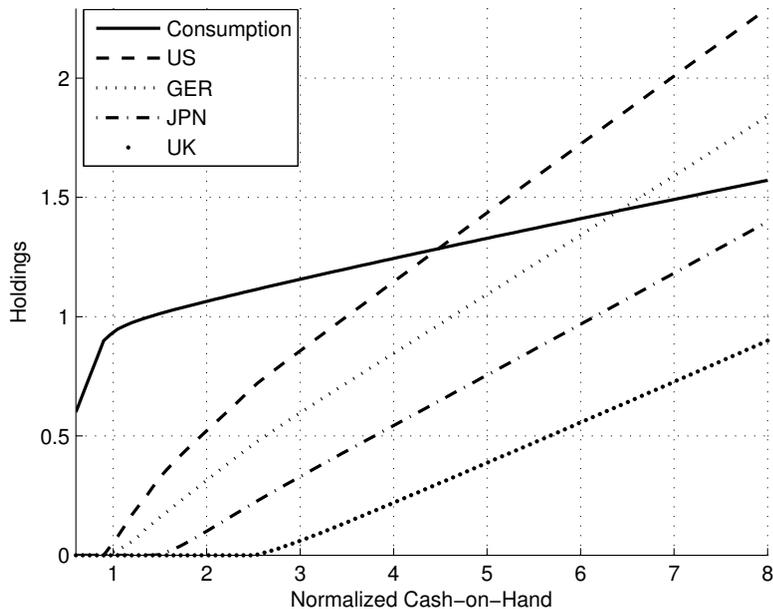

Figure A2: Optimal consumption and investment policy functions as a function of the normalized cash-on-hand.

Optimal policy functions of this counterfactual exercise are reported in Figure A.6. When the agent enters the saving region she starts investing only in the asset with the lowest correlation with aggregate labor income shocks. Only at higher levels of liquid wealth to labor income ratios, the agent starts diversifying her portfolio, adding stocks one at a time. Moreover, the order in which the agents start investing in the different stocks matches the inverse rank of the correlations between labor income and asset returns. As a consequence, the optimal portfolio shares reported in Figure A3 are heavily skewed toward the asset with the lowest correlation.
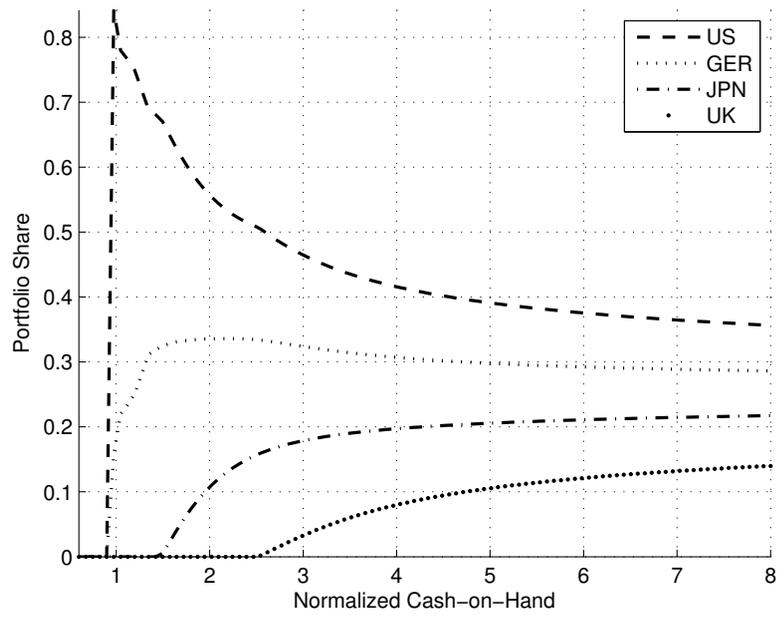
Figure A3: Optimal portfolio shares as a function of normalized cash-on-hand.