Implications of Keeping up with the Joneses Behavior for the Equilibrium Cross Section of Stock Returns: International Evidence*

JUAN-PEDRO GÓMEZ, RICHARD PRIESTLEY, and FERNANDO ZAPATERO

ABSTRACT

This paper tests the cross-sectional implications of "keeping up with the Joneses" (KUJ) preferences in an international setting. When agents have KUJ preferences, in the presence of undiversifiable nonfinancial wealth, both world and domestic risk (the idiosyncratic component of domestic wealth) are priced, and the equilibrium price of risk of the domestic factor is *negative.* We use labor income as a proxy for domestic wealth and find empirical support for these predictions. In terms of explaining the cross-section of stock returns and the size of the pricing errors, the model performs better than alternative international asset pricing models.

JEL Codes: G15, G12, G11.

Keywords: Keeping up with the Joneses, restricted market participation, international asset pricing, local risk, global risk.

---

*Gómez is with the IE Business School, Madrid, Priestley is with the Department of Financial Economics, Norwegian School of Management, and Zapatero is with FBE, Marshall School of Business, USC. Gómez and Zapatero thank the Spanish MICINN for their generous support through research project number ECO2008-02333-RWC. We thank Limei Che and Dmitri Kantsyrev for excellent research assistance. Many helpful comments and suggestions from an anonymous referee are gratefully acknowledged. We also thank Ilan Cooper, Peter DeMarzo, Valery Polkovnichenko, Rosa Rodríguez, Sergei Sarkissian, Tano Santos, Lucie Tepla, Raman Uppal, Robert Stambaugh (the Editor), and Madhu Veeraraghavan for helpful comments. We would especially like to thank John Griffin who generously provided us with the Datastream codes to collect the stock return data in the U.K., Japan, and Germany. The usual caveat applies.

In this paper we test the implications of an international asset pricing model where agents have preferences that are exogenously defined over both their own consumption as well as the contemporaneous average consumption of a reference group, defined in this paper as the agent's countrymen. These preferences are termed "keeping up with the Joneses."

In the absence of frictions, utility based on keeping up with the Joneses preferences yields the symmetric equilibrium described in Galí (1994), in which all agents hold the same global portfolio, and Joneses behavior translates into a lower price of risk on the single systematic risk factor. We consider two different specifications that incorporate frictions. The first is the case of standard keeping up with the Joneses preferences, in which agents' marginal utility of consumption increases with the aggregate consumption of their neighbors (as in Abel (1990) and Galí (1994)), but where some agents do not have access to financial markets. We call this first specification *exogenous* keeping up with the Joneses preferences, since concern for other agents' wealth is directly imposed in the utility function. Our second specification considers a friction in which all agents care about two different goods, a global good and a local good, but the local good is in short supply; in this case, agents optimally take into consideration the wealth of their neighbors because of the effect it might have on the price of the good in limited supply. We call this specification *endogenous* keeping up with the Joneses preferences, since concern about neighbors' wealth arises as an equilibrium result. This model is developed in DeMarzo, Kaniel, and Kremer (2004).

These two specifications —exogenous and endogenous— of keeping up with the Joneses preferences have the following implications in an international setting. First, portfolio holdings of the representative investor of each country differ across countries, as shown in DeMarzo, Kaniel, and Kremer (2004); second, in equilibrium, an approximate multi-beta linear factor model arises where, together with the global systematic risk factor, there is an additional systematic risk factor per country that commands a negative price of risk.

The keeping up with the Joneses effect gives rise to negative prices of risk on the country-specific factors because investors require a premium for holding stocks with no, or negative, correlation with the unhedgeable labor or entrepreneurial income (generally foreign stocks). Symmetrically, investors are willing to pay a premium for those stocks that will have a stronger correlation with the unhedgeable risk of the Joneses (presumably, domestic stocks) since it is these stocks that keep them up with the Joneses. We find empirical support for this prediction, and hence for keeping up with the Joneses behavior in the presence of a constraint on some agents. Note that the model's predictions and the empirical results are in contrast to those that are predicted by models of partial integration, which suggests a positive price of risk on the local risk factor (see Errunza and Losq (1985)).

Our paper differs from DeMarzo, Kaniel, and Kremer (2004) in three main respects. First, we derive the implications of keeping up with the Joneses behavior for the equilibrium cross-section of asset returns. Consistent with the results on portfolio choice in their paper, we show that, qualitatively, the asset pricing predictions are the same whether keeping up with the Joneses behavior is exogenous or endogenous. Second, we test empirically the asset pricing implications of keeping up with the Joneses preferences, whether exogenous or endogenous. Third, although our tests do not allow us to quantify the relative impact of endogenous versus exogenous Joneses, we can assess whether each model in isolation is supported by the data.

To test the asset pricing implications of keeping up with the Joneses preferences we need to proxy for the nondiversifiable income of the reference group (the "Joneses"), whether focusing on the exogenous or endogenous specification. Due to moral hazard and country-specific regulation, human capital (measured by income from labor) is arguably the most obvious nonfinancial, nontradable asset. Moreover, it represents the largest fraction of the individual's personal income across countries. For instance, Baxter and Jermann (1997) observe that, in the U.S., labor's share of national income is about 60%. In a more recent paper, Gollin (2002) corrects the estimates for labor income share in total national income by including the income of the self-employed (like our entrepreneurs). After this correction, labor income share is in the range of 65% to 85% for most countries.

We test the asset pricing implications of both exogenous and endogenous keeping up with the Joneses preferences using stock returns from Germany, Japan, the U.K., and the U.S. We undertake two types of empirical tests. First, we form portfolios of stocks according to size and, following the suggestions in Daniel and Titman (2005) and Lewellen, Nagel, and Shanken (2006), according to size and the lagged loadings of individual stock returns on the measure of country-specific nondiversifiable income. Cross-sectional regressions of these portfolios' average stock returns on the models' factors confirm the predictions that the estimated prices of risk on the country-specific factors are negative. The model can explain a reasonable amount of the cross-sectional variation in international stock returns, more than various other international asset pricing models that use currency risk, macroeconomic risk factors, and the Fama and French (1998) international HML risk factor. Formal analysis of the performance of the model is undertaken by assessing the pricing errors. The pricing errors are smaller than, or similar to, those from other international asset pricing models, lending further support to the role of keeping-up-with-the-Joneses behavior.

The second test is undertaken in order to provide a robustness analysis of the portfolio formation techniques and the two step estimation methodology. Instead of forming portfolios of stock returns, we take randomly chosen individual stock returns from Germany, Japan, the U.K., and the U.S. and estimate the model parameters using nonlinear seemingly unrelated regressions (NLSUR). The simultaneous estimation of betas and prices of risk removes the errors-in-variables problem inherent in the Fama and MacBeth (1973) two-step procedure. Therefore, portfolios do not need to be formed and consequently the critique of asset pricing tests raised by Daniel and Titman (2005) and Lewellen, Nagel, and Shanken (2006) that stems from portfolio formation is no longer a concern.[1] Using this methodology we find further support for the model's predictions.

The empirical results support the joint hypothesis of undiversifiable, nonfinancial income and relative wealth considerations across agents. In a final analysis, given the estimated prices of risk for the (country-specific) risk factors, we back out the implied Joneses parameter for the exogenous model, the market-weighted risk aversion coefficient, and the implied risk aversion coefficient for both the exogenous and endogenous model's representative agent in each country.[2] We perform this exercise for two different values of the relative (as a percentage of the country's total output) undiversifiable, nonfinancial income in each country, one including and one omitting the income of the self-employed. We take the prices of risk from the model estimated using portfolios as well as individual securities. For all countries, the inferred (exogenous) Joneses parameter is very high and lies within the admissible range according to the model (greater than zero and strictly smaller than one). The market-weighted risk aversion coefficient is reasonable using both portfolios and individual securities. Looking at the risk aversion coefficients for the representative agents, in the case of the endogenous Joneses the implied risk-aversion coefficient is always negative, which is clearly inadmissible. In contrast, in equilibrium the implied risk aversion coefficient in the presence of exogenous Joneses behavior takes economically meaningful values for all countries.

The paper is organized as follows. In Section I we discuss the related literature. We introduce the model and derive its testable implications in Section II. Section III presents the data. The main empirical results are reported in Section IV. In Section V we compute the keeping up with the Joneses parameter using our estimated prices of risk. Section VI offers some conclusions.

I. Related Literature

To our knowledge this is the first paper to consider and test the asset pricing implications of keeping up with the Joneses preferences in an international setting. Two prior papers are related to ours. First, Lauterbach and Reisman (2004) also derive a multifactor asset pricing model in the presence of keeping up with the Joneses behavior. However, their model does not explain how these factors arise in equilibrium. More importantly, they do not specify how to test the model nor do they perform any empirical work. Second, Shore and White (2002) study and calibrate an international model with Joneses preferences. They concentrate on portfolio holdings and do not include empirical tests of the asset pricing implications.

Within a purely domestic setting, keeping up with the Joneses has been discussed in Galí (1994) and Abel (1999). “Catching up with the Joneses” preferences (where past, rather than contemporaneous, consumption of the Joneses is the benchmark) have been used in Abel (1990), Ferson and Constantinides (1991), Campbell and Cochrane (1999), Boldrin, Christiano, and Fisher (2001), and Chan and Kogan (2002) as a possible explanation of the equity risk premium puzzle. Chen and Ludvigson (2005) find that a model with habit persistence preference is helpful in explaining the cross-section of U.S. stock returns. Head and Smith (2003) consider Joneses preferences, among others, in their attempt to explain interest rate persistence for a cross-section of countries. Ravina (2005) finds evidence that supports keeping up with the Joneses behavior using micro data in the U.S. Gómez (2007) shows how, in the presence of nondiversifiable income, keeping up with the Joneses behavior may yield biased portfolios in equilibrium.

Our paper is also related to the empirical literature on asset pricing under limited market participation in that we assume the existence of nondiversifiable, nonfinancial income, which may include labor income (see, for example, Campbell (2000) and Haliassos and Michaelides (2002)) and entrepreneurial income (see, for example, Heaton and Lucas (2000) and Polkovnichencko (2004)). Since our primary objective is to study implications of keeping up with the Joneses behavior in the cross-section of stock returns, we isolate the

Joneses effect by assuming, first, all nontradable income is in the hands of the workers, and second, workers cannot invest in any financial asset. In other words, absent any relative wealth effect via endogenous or exogenous keeping up with the Joneses behavior, there is no impact of limited market participation on the equilibrium asset prices.[3]

Regarding the first assumption, it is debatable whether investors with access to financial markets could be endowed with nontradable assets. This would result in an additional hedging demand on the side of investors. Viceira (2001) and Cocco, Gomes, and Maenhout (2005) study this problem. Notice that the effect we would expect from this *direct* hedging demand for nonfinancial income would be the opposite of the *indirect* effect resulting from keeping up with the Joneses preferences: assets positively correlated with the investor's nondiversifiable income would be "crowded out" from the optimal portfolio (whereas, in the case of keeping up with the Joneses preferences, the investor's demand for these assets increases). Hence, investors would demand a higher expected return to hold these assets (contrary to the negative price of risk for the country risk factors resulting from keeping up with the Joneses preferences).

With respect to the second assumption, other models such as Basak and Cuoco (1998), Polkovnichenko (2004), and Gomes and Michaelides (2008) allow the constrained agents to invest in bonds. This simplification could affect our results qualitatively only if bonds provided country-specific risk hedging for the workers nondiversifiable income. However, agents endowed with labor or entrepreneurial income are likely to face short-selling constraints. Moreover, as mentioned in DeMarzo, Kaniel, and Kremer (2004), moral hazard concerns will prevent workers from borrowing against their future nondiversifiable income.

Our results are also related to the literature on international asset pricing. Empirically it appears that domestic asset pricing models are able to price local assets more accurately than international models, and international asset pricing models can be improved upon if domestic factors are also included. For example, Cho, Eun, and Senbet (1986) reject the international Arbitrage Pricing Theory (see also Gultekin, Gultekin, and Penati (1989) and Korajczyk and Viallet (1989)). King, Sentana, and Wadhwani (1994) find that local risk is priced in an international multifactor model. Griffin (2002) claims that the world book-to-market factor is a proxy for a domestic factor. Chan, Karolyi, and Stulz (1992) find support for the role of domestic factors in a conditional version of the International CAPM. Harvey (1991) finds that the international CAPM is rejected for developed markets and Dumas, Harvey, and Ruiz (2003) reject market integration for 12 developed OECD countries. For developed markets such as those considered in this paper, models of partial integration should be becoming less important because various restrictions that investors faced in the past no longer apply. In this paper we identify a specific domestic risk factor that results from keeping up with the Joneses preferences within the framework of complete market integration. The empirical tests show that these preferences have a strong effect on asset prices within an international context. Thus, our finding that local factors are important is consistent with the earlier puzzling rejections of asset pricing models of full market integration even for countries where there are no barriers to trade.

Finally, we mention other asset pricing papers based on nonstandard preferences.

Among the most influential are Campbell and Cochrane (1999), who consider habit formation, but with complete, frictionless markets. Also, Bansal and Yaron (2004) use recursive preferences to develop a model with two factors, aggregate consumption and aggregate wealth, which are priced. Another recent related equilibrium model is DeMarzo, Kaniel, and Kramer (2008), who show how market incompleteness can lead to relative wealth concerns and generate financial bubbles when the investors are willing to pay a premium for securities that align their income with that of their cohort.

II. Implications of Keeping Up with the Joneses Preferences

In this section we study the equilibrium consequences of an economy populated by agents with keeping up with the Joneses preferences. We consider the two main specifications discussed in the literature: exogenous and endogenous keeping up with the Joneses preferences.

In both specifications, we assume a one-period economy with two countries, $k \in \{l, f\}$. In each country there is a local firm. At time $t = 0$, each firm issues one share that will yield a random payoff in time $t = 1$. We normalize the initial value of the firm to one. Let r_k denote the random excess return on a share of firm k. The vector $r = (r_l, r_f)'$ has a joint distribution function $F(r)$. Firm shares in both countries can be freely traded overseas. There is also a risk-free bond in zero net supply. Let R denote the return on the risk-free bond. Financial markets are assumed to be complete.

In each country there are two types of agents, investors and workers.

A. Exogenous Keeping Up with the Joneses Preferences

In this subsection we analyze the implications of a version of the keeping up with the Joneses preferences of Abel (1990) and Galí (1994). In particular, in the economy we consider investors are endowed with utility function of the form[4]

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

where c denotes the investor's consumption of the single consumption good, the economy's numeraire; C is the country average or per capita consumption; $\alpha > 0$ is the (constant) relative risk aversion coefficient; and $1 > \gamma \geq 0$ is the Joneses parameter. For $\gamma > 0$, the constant average consumption elasticity of marginal utility (around the symmetric equilibrium), $\alpha\gamma$, is positive as well: increasing the average consumption per capita C makes the individual's marginal consumption more valuable since it helps her keep up with the Joneses. In short, we assume the country average consumption to be a positive consumption externality. By imposing $\gamma < 1$ we ensure a positive risk aversion coefficient for the representative investor, as will become clear in the next section.[5]

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

Since each investor takes C as exogenous and common, the typical aggregation property of CRRA utility functions allows us to replace all the investors in a given country by a representative investor with utility function (1) endowed with the aggregated investor's income without affecting the equilibrium prices. At time $t = 0$ each representative investor is endowed with a share of the local firm (unit value by assumption); hence, $c^0 = 1$ in both countries.

The representative investor solves the following optimal portfolio problem:

$$\begin{aligned} x^* = \operatorname{argmax}_x \quad & E\, u(c, C) \\ \text{s.t.} \qquad & c = R + r'x, \\ & C = c + w, \end{aligned} \tag{2}$$

where x^* represents the weights invested in firms l and f. The first-order condition from problem (2) yields the usual Euler condition, $E\,(r\, u_c(c, C)) = 0$. Given the utility function (1) and the definitions in (2) we can write the problem's first-order condition as a function of the investor's consumption and the workers' relative wealth, w/c:

$$E\left(r\, c^{-\alpha(1-\gamma)}(1 + w/c)^{\alpha\gamma}\right) = 0. \tag{3}$$

Notice that, in the absence of keeping up with the Joneses behavior ($\gamma = 0$), the previous condition reduces to $E\,(r\, c^{-\alpha}) = 0$, the standard CRRA Euler equation. However, in the presence of keeping up with the Joneses behavior ($\gamma > 0$) we observe two effects:

(i) as in Galí (1994), the risk aversion coefficient for the representative investor is scaled down to $\alpha(1-\gamma)$; and (ii) the marginal utility is "levered" by the relative income of the workers. Due to the keeping up with the Joneses behavior, the more affluent workers become, the higher the marginal utility of consumption for the investor. The impact of this leverage is given by the constant average consumption elasticity of marginal utility (around the symmetric equilibrium), $\alpha\gamma$.

Condition (3) allows us to solve for the representative investor's optimal portfolio. Since financial markets are complete, there exists a mimicking portfolio X^w that maps the workers' relative income onto the investment opportunity set such that $w/c = w^0(R + r'X^w)$. Following Galí (1994), given w^0 and X^w, for small values of $E(r)$, the optimal portfolio of the representative investor of country k can be approximated as a function of α, γ, and the risk-adjusted risk premia $\Omega^{-1}E(r)$, with $E(r)$ and Ω the mean return vector and covariance matrix of r, respectively:[6]

$$x_k^* = \frac{\theta_k \gamma_k}{1-\gamma_k} X_k^w + \frac{1}{\alpha_k(1-\gamma_k)} \Omega^{-1} E(r), \tag{4}$$

where $\theta_k = \frac{w_k^0}{1+w_k^0}$, the workers' initial wealth as a proportion of the country's total wealth (investor's plus nondiversifiable wealth).

Equation (4) shows the expected impact of our assumptions on the investor's portfolio holdings. In the first place, if all agents in the country hold a well-diversified portfolio (that is, if absent any friction, $\theta_k = 0$), the alleged keeping up with the Joneses behavior of investors would translate into a redefinition of the representative agent's risk aversion parameter $\alpha_k(1-\gamma_k)$. The optimal portfolio would be re-scaled accordingly and the only asset pricing implication would be a lower expected market price of risk.

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

B. Endogenous Keeping Up with the Joneses Preferences

In this section, we discuss the endogenous keeping up with the Joneses preferences presented in DeMarzo, Kaniel, and Kremer (2004). In this specification agents consume two types of goods: c, which has the interpretation of a global good, and w_k, $k \in \{l, f\}$, a local good. Utility over consumption for these two goods is given by[7]

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

The parameter $\delta > 0$ specifies the relative importance of the local good. All consumption takes place at the end of the period. At time $t = 0$, investors are endowed with shares of the firm that produces the global good. Call c_k^0 the aggregate value of those shares at the beginning of the period for agents in country k. For simplicity, let $c_k^0 = 1$ in both countries. Workers in each country will receive a fixed number $\bar{w}_k$ of units of the local good at time $t = 1$. In equilibrium, the relative price of the local good in terms of the global good at $t = 1$ is given by $p_k = \delta \left(\frac{c_k}{\bar{w}_k}\right)^{\alpha}$. As would be expected, the scarcer the (fixed) local good endowment relative to the (stochastic) global good consumption, the higher the relative price of the former. The investor's hedging demand for this risk will trigger the endogenous keeping up with the Joneses behavior in this model. Financial markets are complete, as described at the beginning of the section.

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

$$u_c(c,p) = c^{-\alpha}\left(1+\delta^{1/\alpha}p^{1-1/\alpha}\right)^{\alpha}. \tag{5}$$

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

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

With these definitions, and following the same procedure as in Section II.A, we can write the approximate function for the country k investor's optimal portfolio as follows[8]

$$x_k^* = \frac{\theta_k(\alpha_k - 1)}{\alpha_k} X_k^w + \frac{1}{\alpha_k}\Omega^{-1}E(r). \tag{6}$$

Notice that, in this model, the optimal portfolio for the logarithmic investor ($\alpha = 1$) coincides with the benchmark, well-diversified portfolio $\Omega^{-1}E(r)$. No relative wealth concern arises even in the presence of local, nondiversifiable wealth.

C. Equilibrium Asset Pricing Implications

Comparing (4) and (6) we conclude that, whether endogenous or exogenous, the existence of keeping up with the Joneses behavior and nondiversifiable income implies the following optimal portfolio for the representative investor of country k:

$$x_k^* = \theta_k b_k X_k^w + \tau_k \, \Omega^{-1}E(r), \tag{7}$$

where the parameters b and τ represent the portfolio bias and the risk tolerance coefficient, respectively, with values:

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

Market clearing in financial markets at time $t = 0$ requires that $\sum_k c_k^0 x_k^* = (c_l^0, c_f^0)'$. Divide both sides by the global financial wealth $c_l^0 + c_f^0$, call x_M the market portfolio, $x_M = (\omega_l, \omega_f)'$, with $\omega_k = c_k^0/(c_l^0 + c_f^0)$, and let r_M denote the return on the global market portfolio. The market clearing condition then becomes

$$\sum_k \omega_k x_k^* = x_M. \tag{8}$$

Spot market clearing at time $t = 1$ implies that workers consume the proceeds of their (nontradable) endowment, w, and investors the return on their portfolios, c.[9]

Let $\theta_k > 0$ for $k = \{l, f\}$. We regress the workers non-diversifiable wealth return, $r_k^w = r' X^w$, onto the world market portfolio return plus a constant[10]

$$r_k^w = a_k + \beta_k \, r_M + \xi_k. \tag{9}$$

Portfolio $\beta_k \, x_M$ represents the projection of the workers income onto the security market line spanned by the global market portfolio x_M. Define the portfolio $F_k \equiv X_k^w - \beta_k \, x_M$ as a "residual" factor portfolio with return $r_k^F = r' F_k$. By construction, this portfolio has zero covariance with the global market portfolio. Additionally, it has expected excess return $E(r_k^F) = E(a_k) = E(r_k^w) - \beta_k E(r_M)$. The net investment in this portfolio is $(1 - \beta_k)$. For any $\beta_k \neq 1$, let $\frac{F_k}{(1-\beta_k)}$ be the normalized (that is, unit net investment) zero beta portfolio for country k. After these definitions, the workers' portfolio can be represented by the orthogonal decomposition

$$X_k^w = F_k + \beta_k x_M. \tag{10}$$

Equation (10) says that the workers' portfolio can be expressed as a linear combination of the market portfolio and a zero-beta (orthogonal) portfolio. We replace X_k^w in (7) by

(10):

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

This portfolio has three components. Portfolio F_k is country-specific and can be interpreted as a *hedge portfolio*: for each country k, portfolio F_k hedges investors from the risk involved in keeping up with the (domestic) nondiversifiable Joneses risk. Given the orthogonality conditions, this portfolio plays the role of a country-specific, zero-beta asset.

The projection component, $\beta_k x_M$, corresponds to that part of the workers wage income perfectly correlated with the global market portfolio. The standard component, $\Omega^{-1}E(r)$, is the highest global Sharpe ratio portfolio, and it is common across countries.

We define the coefficient H, the inverse of the risk tolerance coefficient, as

$$H^{-1} = \sum_k \omega_k \tau_k.$$

After imposing market clearing (8), we solve for the equilibrium expected returns[11]

$$E(r) = H \, \Omega \left[\left(1 - \sum_k \omega_k \theta_k \, b_k \, \beta_k \right) x_M - \omega_l \theta_l \, b_l \, F_l - \omega_f \theta_f \, b_f \, F_f \right]. \qquad (11)$$

Define the matrix $\boldsymbol{F}$ of dimension $N \times 3$ as the column juxtaposition of the market portfolio and the orthogonal portfolios, $\boldsymbol{F} \equiv (x_M, F_l, F_f)$. Let $\boldsymbol{r^F} \equiv (r_M, r_l^F, r_f^F)$ denote the vector of factors' risk premia. Additionally, define the *wealth vector* $\boldsymbol{W}$ as follows:

$$
\boldsymbol{W} \equiv H \begin{pmatrix} 1 - \sum_k \omega_k \theta_k \, b_k \, \beta_k \\ \\ -\omega_l \theta_l \, b_l \\ \\ -\omega_f \theta_f \, b_f \end{pmatrix}.
$$

Given these definitions, the equilibrium condition (11) can be re-written as follows:

$$
E(r) = \Omega \, \boldsymbol{F} \boldsymbol{W}. \tag{12}
$$

Pre-multiplying both terms of the previous equation by the transpose of matrix $\boldsymbol{F}$, we obtain the equilibrium condition for the vector of prices of risk, $\boldsymbol{\lambda}$, with the market risk premium, λ^M, as the first component:

$$
\boldsymbol{\lambda} = \boldsymbol{F}' \Omega \, \boldsymbol{F} \, \boldsymbol{W}, \tag{13}
$$

where $\boldsymbol{F}' \Omega \, \boldsymbol{F}$ is a matrix of dimension 3×3 whose first column (row) includes the market return volatility and a vector of two zeros and the remaining elements are the covariances between F_l and F_f.

The risk premia on the market and the two zero-beta portfolios is then:

$$
\begin{aligned}
\lambda^M &= H \left(1 - \sum_k \omega_k \theta_k \, b_k \, \beta_k \right) \sigma_M^2, \\
\lambda^l &= -H \left(\omega_l \theta_l \, b_l \, \mathrm{Var}(r_l^F) + \omega_f \theta_f \, b_f \, \mathrm{Cov}(r_l^F, r_f^F) \right), \\
\lambda^f &= -H \left(\omega_l \theta_l \, b_l \, \mathrm{Cov}(r_l^F, r_f^F)) + \omega_f \theta_f \, b_f \, \mathrm{Var}(r_f^F) \right).
\end{aligned} \tag{14}
$$

This system of equations allows us to test the model's predictions. In the first place, the model predicts that all prices of risk should be increasing (in absolute value) in the aggregate risk aversion coefficient H.

The world market portfolio, x_M, is partially correlated with each country's non-diversifiable risk. That correlation, captured by the coefficient β_k, offers partial hedging against deviations from the local Joneses (in case $\theta b > 0$). Therefore, the equilibrium price of risk for the global risk factor, λ^M, varies relative to the symmetric equilibrium: the parentheses in the first equation in (14) (which in the case of a symmetric equilibrium would be one) captures the net price of risk on the global risk factor, after discounting the (capitalization-weighted) Joneses hedging effect. If the weighted value of the betas is higher than the world market beta (i.e., one), the model predicts that the market price of risk could turn negative. Intuitively, if the hedging properties of the market portfolio against Joneses deviations outweigh the compensation for systematic risk, the *net* expected market price of risk becomes negative.

Furthermore, if there is a relative wealth concern ($b > 0$) in the economy (either endogenous or exogenous) and workers' income is not diversifiable ($\theta > 0$), there should be two additional risk factors together with the market risk factor. Regarding their sign, the model predicts that if $\text{cov}(r_l^F, r_f^F) > 0$, then λ^l and λ^f will be negative. To understand this result, suppose for the moment that the zero-beta portfolios were orthogonal ($\text{Cov}(r_l^F, r_f^F) = 0$). Then the price of risk would be easily isolated and strictly negative. The intuition for the negative sign would be as follows: an asset that has positive covariance with portfolio F_k will hedge the investor in country k from the risk of deviating from the nondiversifiable (domestic) income of the Joneses. This investor will be willing to pay a higher price for that asset thus yielding a lower return. In equilibrium, the price of risk for F_k would be, in absolute terms, increasing in b_k and the volatility of the hedge portfolio. If the covariance between both zero-beta portfolios is positive, this just increases the absolute value of the negative prices of risk for every country's hedge portfolio.

Finally, solving for $\boldsymbol{W}$ in (13) and replacing it in (12), we obtain

$$E(r) = \boldsymbol{\beta}\,\boldsymbol{\lambda}, \tag{15}$$

where $\boldsymbol{\beta} = \Omega\,\boldsymbol{F}\,(\boldsymbol{F}'\Omega\,\boldsymbol{F})^{-1}$ denotes the 2×3 (in general $N \times (1+K)$, with N the number of assets and K the number of countries) matrix of betas, with the first column as the market betas for both assets.

We name this pricing model which captures equilibrium implications of keeping up with the Joneses preferences —under both the exogenous and the endogenous specifications— KEEPM, which stands for "KEEping up Pricing Model."

For a given asset $i \in 1, 2, .., N$, the model predicts three betas: the standard (global) market beta and two additional country-specific betas:

$$\begin{pmatrix} \beta_i^l \\ \beta_i^f \end{pmatrix} = \frac{1}{D} \begin{pmatrix} \mathrm{Var}(r_f^F)\mathrm{Cov}(r_i, r_l^F) & - & \mathrm{Cov}(r_l^F, r_f^F)\mathrm{Cov}(r_i, r_f^F) \\ \mathrm{Var}(r_l^F)\mathrm{Cov}(r_i, r_f^F) & - & \mathrm{Cov}(r_l^F, r_f^F)\mathrm{Cov}(r_i, r_l^F) \end{pmatrix},$$

with $D = \mathrm{Var}(r_l^F)\mathrm{Var}(r_f^F) - \mathrm{Cov}^2(r_l^F, r_f^F) > 0$.

To understand the model's prediction in terms of these betas, assume first that both zero-beta portfolios are pairwise orthogonal, $\mathrm{Cov}(r_l^F, r_f^F) = 0$. In this case, an asset positively correlated with country l nondiversifiable Joneses risk ($\mathrm{Cov}(r_i, r_l^F) > 0$) and with no, or negative, correlation with country f nondiversifiable Joneses risk ($\mathrm{Cov}(r_i, r_f^F) \leq 0$) will have $\beta_i^l > 0$ and $\beta_i^f \leq 0$ (the symmetric result follows for an asset i with $\mathrm{Cov}(r_i, r_l^F) \leq 0$ and $\mathrm{Cov}(r_i, r_f^F) > 0$). Notice that if $\mathrm{Cov}(r_l^F, r_f^F) > 0$ (as is the case in our empirical test), the second term in the computation of the first beta is negative and the second term in the computation of the second beta is positive, which (along with the minus sign in front of the second term) goes in the same direction as our previous conclusion about the signs of the betas.

The sign of these betas together with that of the expected price of risk on the orthogonal portfolios in (14) explains the equilibrium expected returns in the KEEPM. Besides the global market risk premium, investors require a premium for holding stocks with no, or negative, correlation with the nonhedgeable local labor or entrepreneurial income (generally foreign stocks). In addition, investors are willing to give up expected returns (that is, pay a premium) for the stocks that are correlated with the idiosyncratic component of the domestic Joneses and therefore help them keep up with the Joneses. This result depends in a fundamental way on the market friction that prevents some agents from participating in the markets. This effect is absent in the analysis of Campbell and Cochrane (1999), who assume complete markets and frictionless markets. Another popular type of model that also postulates nonstandard preferences is the model that uses the recursive preferences of Epstein and Zin (1989) and Weil (1989) (for example, Bansal and Yaron (2004)). These are also consumption-based models in which the pricing kernel has a second factor, besides aggregate current consumption (equivalent to our market portfolio), which is aggregate wealth. This second factor is used to explain the expected return of securities that can hedge the so-called long-run risk (which plays a similar role to our domestic keeping up with the Joneses risk). Finally, DeMarzo, Kaniel, and Kramer (2008) base a model on endogenous relative wealth concerns and show that the resulting negative risk premium can explain financial bubbles in equilibrium.

The rest of the paper deals with testing the KEEPM. In Section IV we analyze the general implications of the KEEPM, both under the endogenous and exogenous specifications, and in Section V we try to distinguish implications of the exogenous and endogenous specifications. Next we turn to a discussion of the data used in the empirical analysis.

III. Data

We test the asset pricing implications of the KEEPM using data from the U.S., U.K., Japan, and Germany. The choice of countries is limited by the availability of a sufficient time series of data that allows for a meaningful test of the model's cross-sectional implications. In tests of domestic asset pricing models it has become standard to sort stocks into portfolios based on firm characteristics such as size and book-to-market. However, both Daniel and Titman (2005) and Lewellen, Nagel, and Shanken (2006) note that testing asset pricing models using firm characteristics to form portfolios can lead to spurious conclusions about the usefulness of a proposed factor. The reason for this is that the factor structure of the portfolios is so strong that any proposed factor that is only weakly correlated with size or book-to-market will be able to price the test assets. That is, testing a new proposed factor on test assets sorted by size and book-to-market is likely to have very low power.

In order to alleviate this concern we follow the recommendations in Daniel and Titman (2005) and Lewellen, Nagel, and Shanken (2006) and augment characteristic-sorted portfolios with portfolios that are sorted by lagged loadings on our proposed factor. The portfolio formation procedure is as follows: Firms are sorted into three size portfolios based on market capitalization (size) as of December of year t. For each individual firm we regress its excess return on the measure of orthogonal income based on the 24 monthly observations leading up to December of year t to obtain the pre-formation factor loading. We then form three subportfolios within each size portfolio based on the pre-formation factor loading. These subportfolio breakpoints are set such that there is an equal number of firms in each of the nine portfolios.[12] We then calculate monthly returns in year $t+1$ for each of these nine (value-weighted) size and factor loading sorted portfolios. The portfolios are re-balanced each year. In addition to these nine portfolios, we also include 10 size sorted portfolios, providing a total of 19 portfolios per country.

The starting period for the analysis is dictated by the availability on Datastream of stock price data for the U.K., Japan, and Germany. To this end we start in January 1973 and collect data up until December 2006. The data from the U.S. comes from CRSP. All returns are in U.S. dollars and excess returns are calculated by subtracting the U.S. one-month T-bill rate from the actual returns.

An additional way of alleviating the problems caused by forming portfolios sorted on firm characteristics highlighted by Daniel and Titman (2005) and Lewellen, Nagel, and Shanken (2006) is to use individual stock returns rather than portfolios. This is impractical in the Fama and MacBeth (1973) methodology due to the errors-in-variables problem that arises from the two-step methodology of estimating betas in the first step and prices of risk in the second step. However, this problem can be circumvented by estimating the betas and prices of risk simultaneously using nonlinear seemingly unrelated regressions. The simultaneous estimation of betas and prices of risk has been used in estimating asset pricing models in, for example, Burmeister and McElroy (1988), Elton, Gruber, and Blake (1995) and Priestley (1997). Furthermore, the use of individual securities is helpful in that the formation of portfolios raises a number of other problems in its own right. Brennan, Chorida, and Subrahmanyam (1998) show that data snooping biases lead to a lack of test power because portfolios are formed on some empirical characteristic found to be relevant in earlier empirical work (Lo and MacKinlay (1990) and Berk (2000)). Roll (1977) shows that portfolio formation may eliminate important return characteristics by averaging into portfolios.

When we estimate the model using NLSUR, we require that the stocks survive the sample period. We also face a constraint on the number of securities (N) that can be used relative to the number of time-series observations (T). In particular, it is required that $NT > K(N + 1)$, where K represents the number of factors, so that the system of equations has more equations than unknowns. In the case of $N > T$, it is problematic to compute the variance-covariance matrix. However, if N is too small, the risk premia will have large standard errors. As a compromise between computational feasibility and a sufficiently large sample so that approximately valid asymptotic inferences can be made, Burmeister and McElroy (1988) suggest 70 stock returns. To this end we include 20 random individual stock returns each from the U.S., U.K., Japan, and Germany, stocks that must have survived the sample period. This gives a total of 80 equations to estimate and is similar to the number used in Burmeister and McElroy (1988) and Priestley (1997).

Rather than simply selecting one random sample of 20 stock returns for each country and estimating the model once, we select 100 random samples of 20 stocks per country. At the beginning of the sample there are 1,507, 303, 658, and 87 stocks that have survived in the entire sample period in, respectively, the U.S., U.K., Japan, and Germany. We select a random sample of twenty stocks from each country. We then place the stocks back into the entire sample for each country. Next, we select a new random sample of stocks for each country. We repeat this procedure 100 times. Clearly, given the small number of sample stocks in Germany the number of truly independent samples is small.

In order to test the model we need to proxy for the nondiversifiable income of the workers in each country. As we mentioned in the presentation of the model, human capital is arguably one of the main sources of nondiversifiable personal income and labor income seems to be the best available proxy for human capital. Therefore, we use the orthogonal (with respect to the world market portfolio) component of each country's labor income series. The data on labor income are wages and salaries in Germany, wages in Japan, personal income in the U.S., and wages in the U.K. This data is collected from Datastream.

In addition to testing the KEEPM directly, we want to compare its performance relative to that of other international asset pricing models. We consider versions of the international CAPM, the international CAPM with currency risk, the international CAPM with an international book-to-market factor, and an international multifactor model that uses international macroeconomic factors. To proxy currency risk we use a trade-weighted index of the U.S. dollar. Other risk factors based on macroeconomic factors are: world unexpected inflation (derived from the IMF world consumer price index), world unexpected industrial production (derived from the OECD aggregate industrial production index), and the return on world government bond markets (derived from the Lehman Brothers world government bond market index). The unexpected inflation and industrial production factors are the residuals from autoregressions whilst all other factors are return-based. All the data used in the construction of the risk factors are collected from Datastream except for the HML factor, which is kindly provided by Ken French.

IV. Empirical Results

Using portfolio data the model can be consistently estimated by the cross-sectional methods of Fama and MacBeth (1973). We consider the performance of the KEEPM —encompassing both exogenous and endogenous specifications— in terms of a number of metrics. First, and most importantly, we are interested in assessing whether the model's risk factors have the correct sign. A second issue is how well the model does in terms of describing the cross-sectional variation in international stock returns.

We report the adjusted R^2, $\overline{R}^2$, of the cross-sectional regression, which calculates the amount of cross-sectional variation that is captured by the model. Following Jagannathan and Wang (1996) and Lettau and Ludvigson (2001) we calculate the R^2 as $\left[Var_c(\overline{r}_i) - Var_c(\overline{e}_i)\right]/Var_c(\overline{r}_i)$ where Var_c is the cross-sectional variance, $\overline{r}_i$ is the average return, and $\overline{e}_i$ is the average residual. This metric provides a general indication of the fit of the model but it is not a direct test of the model. For that we examine whether the models' pricing errors are zero. Within the above testing framework we also look at how well other international asset pricing models perform.

Using the portfolio returns for the four countries, equation (15) implies a five-factor model with the world market price of risk, the U.S. orthogonal labor price of risk, the U.K. orthogonal labor price of risk, the Japanese orthogonal labor price of risk, and the German orthogonal labor price of risk:

$$E(r_{i,t}) = \lambda^M \beta_i^M + \lambda^{ous} \beta_i^{ous} + \lambda^{ouk} \beta_i^{ouk} + \lambda^{ojp} \beta_i^{ojp} + \lambda^{ogm} \beta_i^{ogm},$$

where $E(r_{i,t})$ is the expected excess return on asset $i \in 1, ..., N$ at time $t \in 1, ..., T$, β_i^w is stock i's β with respect to the world stock market portfolio, λ^M is the world stock market price of risk, β_i^{ous} is stock i's β with respect to the orthogonalized U.S. labor return, λ^{ous} is the U.S. orthogonalized labor price of risk, β_i^{ouk} is stock i's β with respect to the orthogonalized U.K. labor return, λ^{ouk} is the U.K. orthogonalized labor price of risk, β_i^{ojp} is stock i's β with respect to the orthogonalized Japanese labor return, λ^{ojp} is the Japanese orthogonalized labor price of risk, β_i^{ogm} is stock i's β with respect to the orthogonalized German labor return, and λ^{ogm} is the German orthogonalized labor price of risk. The model predicts that $\lambda^{ous} < 0$, $\lambda^{ouk} < 0$, $\lambda^{ojp} < 0$, and $\lambda^{ojp} < 0$.

The Fama and MacBeth (1973) procedure involves a first step in which time-series regressions are used to estimate the betas, and a second step in which cross-sectional regressions are used to estimate the lambdas. Because we have a relatively short sample and have already lost 24 observations in the portfolio formation procedure, we employ a rolling regression approach that uses 36 observations up to time t in the first step to obtain the first beta. We then use this beta in the second step to estimate a cross-sectional regression of average returns at time $t + 1$ on the beta estimated up until time t. The data are then rolled forward one month and the procedure is repeated. This results in a time series of cross-section estimates of the relevant price of risk. In total we lose 60 observations given the portfolio formation and rolling beta estimation method. Therefore, the first cross-sectional estimation period is January 1978, providing a total of 29 years.

The main empirical results of the paper are presented in the first row of Table I, where we report estimates of the KEEPM using 76 portfolios, 19 from each of the U.S., U.K., Japan, and Germany. The orthogonal labor prices of risk of the U.K. and Germany are estimated to be negative, as the model predicts, and statistically significant. They are also economically meaningful, with values of -0.218% and -0.537% per month, respectively. Overall, the model appears to have empirical support.

[INSERT TABLE I]

Both Daniel and Titman (2005) and Lewellen, Nagel, and Shanken (2006) note that interpreting the success of a factor model based on the cross-sectional $\overline{R}^2$ can be dangerous when using size and book-to-market portfolios. This is because obtaining a high $\overline{R}^2$ is not difficult for any factor that is weakly correlated with size or book-to-market. However, simulations in Lewellen, Nagel, and Shanken (2006) show that it is much harder to get a reasonable $\overline{R}^2$ when size and book-to-market portfolios are augmented with portfolios not related to these factors, such as the lagged factor loadings that we use. Therefore, in our case, examination and comparisons of $\overline{R}^2$ should be meaningful. The KEEPM explains 31% of the cross-sectional variation in excess returns, which is reasonable when we consider that we use portfolios of stock returns from four different countries. Overall, the empirical results provide evidence consistent with the theoretical model's predictions.

There are not many other data series that could proxy for personal nondiversifiable income. Housing is a potential candidate (for example, it might capture a part of the entrepreneurial income that we discuss in the model section) and thus we use it as a second measure of orthogonal local personal income for robustness purposes. For the U.K. we use the Halifax Building Society's house price index. The U.S. housing data comprise the median price of sales of existing single-family homes. Monthly data on house price indices are not available for Japan and Germany. The results using housing data, which are only available from 1983 due to lack of housing data back further in time in the U.K., indicate that the orthogonal prices of risk in both the U.K. and the U.S. are negative and statistically significant. The actual estimates (t-statistics) are -0.808 (2.88) and -0.585 (2.64) for the U.S. and U.K., respectively. Thus, we find further support for the model's predictions when using an alternative measure of local orthogonal wealth.

Whilst our central concern is to test our theoretical model, we also consider its performance and robustness relative to a class of other international asset pricing models. The first model is the international CAPM (ICAPM); see Black (1974). This model assumes standard preferences and complete integration of capital markets. It also assumes that purchasing power parity (PPP) holds. Thus, we have:

$$E(r_{i,t}) = \lambda^{ICAPM} \beta_i^{ICAPM},$$

where λ^{ICAPM} is the ICAPM market price of risk and β_i^{ICAPM} is stock $i's$ β with respect to the excess return on the world stock market portfolio. We also consider the model of Fama and French (1998), who suggest a two-factor model for international asset pricing that includes the excess return on the world stock market portfolio and the international high minus low book-to-market factor (HML):

$$E(r_{i,t}) = \lambda^{M} \beta_i^{M} + \lambda^{HML} \beta_i^{HML},$$

where λ^{HML} is the price of risk associated with the HML risk factor and β_i^{HML} is the β with respect to the HML risk factor.

Since it is well known that PPP does not hold, at least in the short and medium terms (see, for example, Grilli and Kaminsky (1991), Wu (1996), and Papell (1997)), investors may be exposed to real exchange rate risk. Theoretical models that incorporate currency risk include Solnik (1974), Stulz (1981), and Adler and Dumas (1983). The empirical specification that we examine is

$$E(r_{i,t}) = \lambda^M \beta_i^M + \lambda^{cb} \beta_i^{cb},$$

where λ^{cb} is the currency index price of risk and β_i^{cb} is the β with respect to the currency index.

In addition to exchange rates, other macroeconomic factors have been used in international asset pricing models (see, for example, Ferson and Harvey (1994)). Therefore, we also examine a model that includes the currency basket and three macroeconomic-based factors: world unexpected inflation, world unexpected industrial production, and the return on world money markets:

$$E(r_{i,t}) = \lambda^{ui} \beta_i^{ui} + \lambda^{uip} \beta_i^{uip} + \lambda^{gb} \beta_i^{gb} + \lambda^{cb} \beta_i^{cb},$$

where λ^{ui} is the inflation price of risk, β_i^{ui} is the β with respect to unexpected inflation, λ^{uip} is the industrial production price of risk, β_i^{uip} is the β with respect to unexpected industrial production, λ^{gb} is the world government bond market price of risk, and β_i^{gb} is the β with respect to the return on the world government bond market.

The rest of Table I reports the estimates of the various models discussed above. In the second row an estimate of the international CAPM is presented. According to the $\overline{R}^2$, which is 4%, the ICAPM, like its domestic counterpart, is unable to explain much of the cross-sectional variation in average returns, and as we will see later, the pricing errors from this model are large. This is interesting since it suggests that most of the explanatory power in the KEEPM model is coming from the local factors and not the market factor.

The next model, in row 3, includes the world HML factor and the world stock market factor. The estimate of the HML price of risk is not statistically significant. Including the HML factor has little impact on the estimate of the world stock market factor. This model can explain 20% of the cross-sectional variation in the asset returns.

The fourth row of Table I reports the model that includes a measure of currency risk along with the world stock market portfolio. The estimated price of currency risk is -0.421% per month and is marginally statistically significant. This model explains 21% of the cross-sectional variation in the excess stock returns. The final row of the table provides results from a version of an international multifactor model that uses only macroeconomic-based risk factors. In this case only the currency risk factor is statistically significant and the $\overline{R}^2$ shows that 21% of the variation in average returns is captured by the model. In summary, Table I shows that the KEEPM model has empirical support and that it performs reasonably well when compared to alternative international asset pricing models that have been used in the extant literature.

A. Individual Security Returns

In this part of the paper we examine the robustness of the results presented so far by analyzing individual security returns instead of portfolios. Due to the use of individual securities we employ a NLSUR methodology that, recall, has the advantage over the traditional Fama and MacBeth (1973) two-step methodology that it avoids the errors-in-variables problem of estimating betas in one step and then the prices of the risk in a second step. Moreover, using NLSUR allows for correlations in the residual variance-covariance matrix that lead to more efficient estimates (both asymptotically and in most small samples; see Shaken and Zhou (2006)).[13]

The drawback of using individual securities is that the firms have to survive the sample period. This induces some survivorship bias in the sample. However, in a relatively short sample such as the one used here the extent of survivorship bias is likely to be limited and is unlikely to effect the signs of the estimated risk premia.

Given a k factor model and a set of N test assets over T observations, the asset pricing model can be expressed as

$$\mathbf{r}_t = E(\mathbf{r}) + \beta_k \mathbf{f}_{kt} + \mathbf{u}_t \tag{16}$$

$$E(\mathbf{r}) = \beta_k \lambda_k, \tag{17}$$

where $\mathbf{r}_t$ is a N vector of excess security returns, $\mathbf{f}_{kt}$ is a k vector of observations on the k risk factors, β_k is a $N \times k$ matrix of betas (sensitivities of returns to the factors), $\mathbf{u}_t$ is a N vector of residual error terms, $E(\mathbf{r})$ is a N vector of expected excess returns and λ_k is a k vector of prices of risk. Substituting equation (17) into (16) and stacking the equations for the N securities gives

$$\mathbf{r} = \{\mathbf{I}_N \otimes [(\lambda' \otimes \iota_{\mathbf{T}}) + \mathbf{f}]\} \beta + \mathbf{u}, \tag{18}$$

where $\mathbf{r}$ is a $NT \times 1$ vector of excess returns, λ is a $k \times 1$ vector of prices of risk, $\mathbf{f}$ is a $T \times k$ matrix of observations of the k factors, β is a $Nk \times 1$ vector of sensitivities, $\mathbf{I}_N$ is a $N \times N$ identity matrix, and $\otimes$ is the Kronecker product operator. The NLSUR estimators are those that solve the following minimization problem:

$$\min_{\lambda,\beta} \mathbf{u}' \left(\hat{\boldsymbol{\Sigma}}_{\mathbf{u}}^{-1} \otimes \mathbf{I}_T \right) \mathbf{u}, \tag{19}$$

where $\hat{\boldsymbol{\Sigma}}_{\mathbf{u}}^{-1}$ is the residual covariance matrix obtained from estimating (18).

[INSERT TABLE II]

Recall from the data section that we have chosen 100 random samples of 20 stocks from each of the four countries. We estimate the model 100 times and report the means, standard deviations, and the 75th and 25th percentiles for both the estimates and t-statistics (in parentheses). Table II reports the findings and shows that the mean estimate of all four prices of risk on the orthogonal local labor income measures are negative and statistically significant with the exception of Japan. Except for Germany, the mean estimates are larger for the individual securities than for the portfolios. Even at the 75th percentile the estimates for the U.S. and U.K. are negative and highly statistically significant. Interestingly, when using individual securities we find that the world stock market price of risk is estimated to be positive and is reasonable at 0.574% per month, which translates into an annual return similar to the actual market premium. The mean $\overline{R}^2$ is 51%, which indicates that the model does quite well in describing the cross-sectional variation in individual asset returns in the four countries. At the 25th percentile the $\overline{R}^2$ is 30% and at the 75th percentile it is 71%. The findings in Panel A of Table II support the evidence in Table I, which suggests that keeping up with the Joneses behavior affects asset prices.

The remaining panels of Table II report estimates from the alternative international asset pricing models considered in Table I. Panel B provides the distribution of the estimates and t-statistics for the ICAPM. Although the estimated market price of risk is positive and statistically significant, and the standard deviation of the estimate is small, the model cannot explain any of the cross-sectional variation in returns. Even at the 75th percentile the $\overline{R}^2$ is only 6%. Panel C reports the results using the FF HML factor, which is estimated to be statistically significant, but has a negative sign. Although the estimate is significant the explanatory power of the model is poor with a mean $\overline{R}^2$ of only 3%. The results from estimating the ICAPM with currency risk are reported in Panel D and indicate that whilst the estimated sign and size of the coefficient on currency risk are similar to those obtained in Table I when using portfolio data, the mean t-statistic indicates that the coefficient is not statistically significant on average. Whilst the estimate is statistically significant for the 25th percentile, at the 75th percentile the estimate is insignificant. The $\overline{R}^2$ is 12% for the individual securities, less than that obtained using portfolios.

Panel E reports the results from the model that uses macroeconomic factors. According to the mean t-statistics none of the estimates are statistically significant. However, the model explains on average 69% of the cross-sectional variation in returns. The reason for this is that the standard deviation of the estimates are large and the actual estimates switch from significant and positive to significant and negative. Unlike the KEEPM model where the estimates are consistently negative across the use of portfolios or individual securities, the signs on the currency factor and the inflation factor have different signs when using portfolios as compared to individual securities.

Overall, the findings in Panel A of Table II provide further support for the KEEPM model. Based on the level of statistical significance and the $\overline{R}^2$ the results are stronger here than in Table I when using portfolios, although they are qualitatively the same. The reason that the prices of risk are more statistically significant in the case of individual securities is likely to stem from the fact that the NLSUR method provides more efficient estimates and the omission of small stocks when using individual securities.

B. Analysis of Pricing Errors

Table III reports the estimates of the pricing errors across aggregated portfolios. Panel A reports the square root of the average squared pricing errors for nine aggregated portfolios. For each of the models in Table I, we report pricing errors by portfolio formation method (only Size, S, and Size and Factor Loadings, SFL) for each country and across all 76 portfolios. Standard errors are reported in parentheses. The column headed KEEPM shows that the pricing errors across all the portfolios are 0.29%. The next column reports the pricing errors from the ICAPM. For most portfolios and countries the pricing errors are smaller in the KEEPM than the ICAPM. Similar results obtain for the models that include the HML and currency basket factors (columns 4 and 5). The model that includes the macroeconomic factors has slightly higher pricing errors across all portfolios than the KEEPM. The patterns in the pricing errors across countries for each model are similar to those of the pricing errors for the aggregated portfolios.

[INSERT TABLE III]

In Panel B we test whether the average pricing errors are significantly different from zero using an asymptotic χ^2 (Wald) test of the null hypothesis that all the pricing errors are jointly zero. In the majority of cases we reject the null hypothesis that the pricing errors are jointly zero, and in every case we reject the null hypothesis when we test across all pricing errors (rather than a group of them). Only in the case of Germany is the null hypothesis of zero pricing errors not rejected at the 5% level for some of the models.

So far, we have established that the pricing errors for the KEEPM are smaller than those for other models, but in most cases it is possible to reject the null hypothesis that the pricing errors are zero. The final row of the table tests whether the pricing errors from the KEEPM are smaller than those from the other models. Under the column ICAPM we show that we can reject the null hypothesis that the pricing errors from the KEEPM are the same as those of the ICAPM. Whilst the pricing errors from the KEEPM are smaller than those from the remaining models, it is not possible to reject the null hypothesis that the pricing errors are the same.

Panel A of Table IV reports the square root of the average squared pricing errors for the five estimated versions of the model using individual securities. The pricing errors for the individual securities are the average pricing errors across all individual assets in all four countries. We then take the average of these across the 100 versions of the model that we estimate. We report the pricing errors, along with standard errors in parentheses, aggregated across all stocks. The pricing errors for the individual securities are similar in size to those of the portfolios. The pricing errors from the KEEPM are 0.236%, which is considerably lower than the pricing errors from the alternative models except the model that uses macroeconomic factors, which has a pricing error of 0.21%, slightly lower than the KEEPM model.

[INSERT TABLE IV]

Tests of the null hypothesis that the pricing errors are jointly zero are reported in Panel B. The null hypothesis that the pricing errors are jointly zero is rejected in every model, similar to the case of the portfolios. We reject the null hypothesis that the pricing errors from the KEEPM model are the same as those from the other models. In particular, we find that the pricing errors from the KEEPM are statistically lower than those of the other models with the exception of the model with macroeconomic factors, which has statistically lower pricing errors than the KEEPM.

In summary, the analysis of the model's pricing errors provides further empirical support for the KEEPM, under both the exogenous and the endogenous specifications, by showing that the pricing errors are smaller for the KEEPM than for most other models. In general, we reject the null hypothesis that the pricing errors are jointly zero across most portfolios, across pricing errors within a given country and type of portfolio formation methodology, and across individual securities. However, the pricing errors from the KEEPM model are lower than those from other models when testing using portfolio returns and are lower than those from all models except the IMACRO model when using individual stock returns.

V. Inference on the Model Parameters

In this section of the paper we attempt to assess whether the inferred parameters of the model are consistent with an exogenous specification, motivated by *status comparison*, or a model where the endogenous *community effect* arises by the relative price of non-tradable assets in limited supply. It is important to stress here that both explanations are not mutually exclusive and, as DeMarzo, Kaniel, and Kremer (2004) state, "it is possible that both status and community effects are important. In this case the two effects will reinforce each other, increasing the tendency for investor herding."

Ideally, we would like to estimate the parameters of the model directly via, for example, maximum likelihood or GMM. However, it is not possible to identify them individually.[14] Therefore, we propose to back out the deeper parameters of the model using the estimated prices of risk. Whilst this approach is limited in that we cannot obtain standard errors, it should provide a reasonable estimate of the direction of the effect and therefore sheds light on the underlying economic mechanism at work.

Given that there are k estimated prices of risk, it is possible from (14) to solve for H and b_k as a function of $\{\lambda^k, \omega_k, \theta_k, \beta_k\}$, the global market portfolio volatility, σ_M^2, and the covariance matrix of the orthogonal portfolios. We perform these exercises using the prices of risk estimated from the portfolios. As a robustness test, we repeat the same exercise using the prices of risk estimated with individual securities as reported in Panel A of Table III.

As a starting point we focus on the prices of risk presented in the first row of Table

I. However, notice that the market price of risk is insignificant. In order to obtain a meaningful estimate for the aggregate risk aversion coefficient H, we impose an exogenous (monthly) value of $\lambda_M = 0.43$, which is simply the mean monthly average excess return on the world stock market portfolio. We then take the residuals from this model and re-estimate the Joneses price of risk on the orthogonal risk factors.[15] The new estimates are -0.034% for the U.S., -0.185% for the U.K., -0.224% for Japan, and -0.627% for Germany, which are very close to those presented in the unconstrained model in row 1 of Table I. These are the estimates that we will use, along with the positive market price of risk of 0.43%, to provide values for the model's deeper parameters in Panel B. For robustness, we re-estimate the parameters for a lower value of $\lambda_M = 0.215\%$ (Panel A), which is 50% lower than the mean market premium, and a higher value of $\lambda_M = 0.645\%$ (Panel C), which is 50% higher than the mean. The estimates of the prices of risk remain virtually constant across the three panels.[16]

The values for the betas are $(\beta_{US}, \beta_{UK}, \beta_{JP}, \beta_{GM}) = (0.0102, -0.0156, -0.0107, -0.0095)$, estimated over the period January 1975 through to December 2006. The global stock market return variance is 0.1684% per month. The weights of each country in total market value are normalized to sum up to one: $(\omega_{US}, \omega_{UK}, \omega_{JP}, \omega_{GM}) = (64.50\%, 13.22\%, 16.88\%, 5.40\%)$, as reported by Morgan Stanley International in 2004.

For the values of θ, the proportion of local nondiversifiable income over total country income, we take the shares reported in Gollin (2002), Table II. For every country we take two possible values: the "naive calculation," defined as employed compensation relative to national output, and the "adjusted calculation," which includes the income of the self-employed as labor income. Given the interpretation of θ as nondiversifiable income, the adjusted calculation is probably more appropriate. Gollin's paper does not report any values for Germany so we assume the the same values for Germany as for the U.K.

[INSERT TABLE V]

For each country and for the two values of θ under consideration, Table V reports the inferred estimates of the portfolio bias b, the Joneses parameter γ, and the relative risk aversion coefficient for the representative agent in the endogenous model, α_{EN}. The estimated aggregate risk aversion coefficient is equal to $H = 0.6683$ in Panel A, $H = 1.9640$ in Panel B, and $H = 3.2299$ in Panel C, all reasonable values. Of course, as expected, the higher the assumed market's risk premium, the higher the implied aggregate risk aversion coefficient.

Given the definition of b in Subsection II.C, we can interpret this parameter as a function of the exogenous keeping up with the Joneses preferences parameter γ, or, in the case of the endogenous model, the investor's risk aversion coefficient, α_{EN}. In the exogenous case, keeping up with the Joneses behavior arises for values of $0 < \gamma < 1$. The endogenous keeping up with the Joneses effect arises for $\alpha_{EN} > 1$.

The estimates for the parameter γ are, for all countries and across the three panels, within the admissible range and very high. This is consistent with strong exogenous keeping up with the Joneses behavior. As for the estimates of the parameter α_{EN}, they are negative for all countries and θ values, which is inconsistent with the predictions of the endogenous model.

We calculate the implied risk aversion coefficient in the absence of keeping up with the Joneses behavior, α_{EX}, and the scaled risk aversion coefficient under keeping up with the Joneses behavior, $\tau^{-1} = \alpha_{EX}(1-\gamma)$. They are reported in Table V, columns 5 and 6, respectively. For this purpose we take the equilibrium definition $H^{-1} = \sum_k \omega_k \tau_k$ and assume that α_{EX} is common across countries. Given the definitions of b and τ in the exogenous relative wealth concerns model we obtain $\alpha_{EX} = H \sum_k \omega_k (1 + b_k)$. The implied value of the risk aversion coefficient for the naive (i.e., lower) estimates of θ is $\alpha_{EX} = 100.67$ in Panel A, $\alpha_{EX} = 100.80$ in Panel B, and $\alpha_{EX} = 100.92$ in Panel C. For the adjusted (i.e., higher) values of θ, $\alpha_{EX} = 72.15$ (Panel A), $\alpha_{EX} = 72.72$ (Panel B), and $\alpha_{EX} = 73.28$ (Panel C). In either case, α_{EX} has implausibly high values. This result is probably due to the low standard deviation of labor income during the period under consideration, especially in the U.S. (7.42% annual) and the U.K. (12% annual) compared to Japan and Germany where the standard deviations are 23% and 20% per annum, respectively.

In the presence of keeping up with the Joneses behavior, however, our exogenous parameter specification model predicts that the representative agent's risk aversion coefficient, τ^{-1}, will be scaled down by $(1-\gamma)$ relative to α_{EX}. The final column of Table V reports the estimates of τ^{-1} and shows that for both the lower and higher values of θ and across panels, the estimates are economically meaningful. In contrast, the estimated risk aversion coefficients in the endogenous relative wealth concern case are negative for all countries and values of θ.

[INSERT TABLE VI]

Table VI backs out the model's deeper parameters using individual securities. Panel A reports the inferred parameters when using the mean lambdas from Panel A of Table II. The market-weighted risk aversion coefficient becomes $H = 2.1672$. The exogenously specified keeping up with the Joneses parameter and the endogenous risk aversion coefficient are similar to the values reported for portfolios in Table V. The estimates of the implied risk aversion coefficient for the exogenous model are $\alpha_{EX} = 510,929$ for the lower values of θ and $\alpha_{EX} = 379$ for the higher values of θ. Both are higher than in the case of portfolios. After taking into account the keeping up with the Joneses parameter γ, the implied risk aversion coefficients for the representative investor are reported in the last column of Table VI. All the estimates are reasonable with the exception of Japan, where they are arguably too high (25.42 and 24.21 for the naive and adjusted θs, respectively).

For the sake of comparison, Panel B reports the inferred parameters after imposing the ad hoc market price of risk $\lambda_M = 0.43\%$ as in Table V, Panel B. We then take the residuals from this model and re-estimate the prices of risk on the orthogonal risk factors (which are very similar to the original prices of risk reported in Panel A). The aggregate risk aversion coefficient turns out to be $H = 1.3135$. The estimates of the implied risk aversion coefficient for the exogenous model are $\alpha_{EX} = 510.30$ for the lower values of θ and $\alpha_{EX} = 378.33$ for the higher values of θ. Both are higher than in the case of portfolios (Table V) and virtually identical to the values in Table VI, Panel A. Comparing Panel B in tables V and VI we observe that the estimate of b becomes much larger in the case of individual securities for the U.S. and the U.K. The Joneses parameter γ is large and below one in both cases; the endogenous risk aversion coefficient α_{EN} is negative in both tables. Relative to the estimates in Table V, the implied Joneses adjusted risk aversion coefficient, τ^{-1}, in Table VI is higher for Japan and Germany, lower for the U.S., and very similar for the U.K.

Overall, the empirical evidence suggests that the price-driven (endogenous) explanation of the keeping up with the Joneses preferences cannot generate the expected prices of risk predicted by the model and confirmed by the data. Although the estimated parameters are consistent with an explanation fully based on the exogenous preferences used in this paper, two caveats should be mentioned here. First, by definition, H is an average across *all* countries in the world. The U.S., U.K., Japan, and Germany together represent about 65% of the world market capitalization. Assuming that the orthogonal prices of risk for the remaining countries are also negative, our estimate of H should be considered a lower bound. Second, our tests cannot rule out the (most plausible) scenario in which the endogenous and exogenous effects reinforce each other.

VI. Conclusion

In this paper, we derive and test in an international setting equilibrium pricing implications of "keeping up with the Joneses" preferences with frictions. We consider two specifications present in the literature, which we call exogenous and endogenous keeping up with the Joneses preferences depending on whether the concern about the wealth of a reference group is directly imposed in the utility function (as in Abel (1990) and Galí (1994)) or the result of competition over a good in short supply (as in the model of DeMarzo, Kaniel, and Kremer (2004)). Equilibrium prices for both the exogenous and endogenous versions are explained by a multifactor model where both world and domestic risk (the nondiversifiable component of domestic wealth) are priced. The equilibrium price of risk associated with domestic risk is negative, since investors are willing to pay more for assets that help them to keep up with the local undiversifiable wealth component. This is in contrast to the positive price of risk that arises in models of partial integration.

We test the model using securities from the U.S., U.K., Japan, and Germany, and the time series of domestic labor income as a proxy for undiversifiable domestic wealth. Using both portfolios and individual securities, and using alternative estimation methodologies, we find that the domestic risk factor has a negative risk premium as predicted by the equilibrium implications of keeping up with the Joneses preferences. When we study the pricing errors, we show that keeping up with the Joneses preferences do a better job explaining security prices than a number of well-established international asset pricing models. We also use data on housing (instead of labor income) as a robustness test and obtain similar conclusions.

Using our risk premium estimates, we back out the parameters of each of the two specifications (exogenous and endogenous) considered in the paper. The derived values are consistent with the exogenous keeping up with the Joneses parametrization, but are out of the acceptable range for the endogenous wealth concerns specification. The derived risk aversion coefficient for the representative agent is only acceptable after accounting for the risk adjustment implied by keeping up with the Joneses. However, we cannot rule out the plausible scenario that both effects are simultaneously in play and reinforcing each other.

Appendix

Let X_1, X_2 be a set of regressors (independent variables) and y the dependent variable. Here, we set $X_1 = r_M$, the return on the world market portfolio; $X_2 = (r_l^w, r_f^w)$, the return on local (l) and foreign (f) nondiversifiable wealth; and $y = r_i$, the return on a given asset/portfolio i. Consider now the following three time-series regressions:

Regression 1: $y = X_1\beta + \epsilon$. Call $\hat{\beta} = (X_1'X_1)^{-1}X_1y$ the OLS estimator of the slope in this regression. This would be the regression of the ICAPM with "no frictions," that is, where the world market portfolio is the only risk factor.

Regression 2: $y = X_1\beta_1 + X_2\beta_2 + \epsilon$. Call

$$\begin{aligned}\hat{\beta}_1 &= (X_1'X_1)^{-1}X_1(y - X_2\hat{\beta}_2)\\ \hat{\beta}_2 &= (X_2'M_1X_2)^{-1}X_2'M_1y\end{aligned}$$

the corresponding slope estimates, with $M_1 = (I - X_1(X_1'X_1)^{-1}X_1)$. This regression would correspond with the equilibrium test of keeping up with the Joneses preferences (after imposing market clearing) *but without orthogonalizing* the local, nondiversifiable wealth returns.

Call $\hat{\gamma}_2 = (X_1'X_1)^{-1}X_1X_2$ and $e_2 = M_1X_2$ the slope estimate and the residuals, respectively, of the regression of X_2 on X_1. Then, the orthogonal decomposition in (9) can be written as

$$X_2 = X_1\hat{\gamma}_2 + e_2.$$

Regression 3: $y = X_1 b_1 + e_2 b_2 + \epsilon$, with $X_1 \perp e_2$ by construction. This is the time-series regression in (12). It is easy to show (based on the Frisch-Waugh Theorem in Greene (2007), page 28) that $\hat{b}_1 = \hat{\beta}$ and $\hat{b}_2 = \hat{\beta}_2$.

From the previous definitions, it follows that

$$\hat{\beta}_1 = \hat{b}_1 - \hat{\gamma}_2 \hat{b}_2. \tag{A1}$$

Let's look now at the cross-sectional regressions:

Cross-section regression 1:

$$y = \lambda_1 \hat{\beta}_1 + \lambda_2 \hat{\beta}_2 + \epsilon. \tag{A2}$$

This is the cross-sectional regression *without orthogonalizing* local nondiversifiable wealth returns.

Cross-section regression 2:

$$y = \lambda_1' \hat{b}_1 + \lambda_2' \hat{b}_2 + \epsilon', \tag{A3}$$

which corresponds to our cross-sectional test (15). Replacing (A1) for $\hat{\beta}_1$ in (A2), the latter can be written as follows:

$$y = \lambda_1 \hat{b}_1 + (\lambda_2 - \hat{\gamma}_2 \lambda_1) \hat{b}_2 + \epsilon. \tag{A4}$$

Comparing equations (A3) and (A4), we have that the prices of risk from equation (14) satisfy

$$
\begin{aligned}
\hat{\lambda}_1' &= \hat{\lambda}_1, \\
\hat{\lambda}_2' &= \hat{\lambda}_2 - \hat{\gamma}_2 \hat{\lambda}_1.
\end{aligned}
$$

This equation shows us that (i) for the global market price of risk, the orthogonalization makes no difference ($\hat{\lambda}_1'$ and $\hat{\lambda}_1$ coincide); and (ii) the price of risk on the country-specific orthogonal residuals, $\hat{\lambda}_2'$, yields the return on the local wealth *net* of diversifiable local risk reward, $\hat{\gamma}_2 \hat{\lambda}_1$, perfectly correlated with the global market return. The KEEPM, either exogenous or endogenous, predicts that this net price of risk should be negative: it hedges investors against the local, *nondiversifiable* risk.

Appendix

Abel, Andrew, 1990, Asset prices under habit formation and catching up with the Joneses, *American Economic Review* 80, 38-42.

Abel, Andrew, 1999, Risk Premia and term premia in general equilibrium, *Journal of Monetary Economics* 43, 3-33.

Adler, Michael, and Bernard Dumas, 1983, International portfolio choice and corporation finance: A synthesis, *Journal of Finance* 38, 925-984.

Bansal, Ravi, and Amir Yaron, 2004, Risks for the long run: A potential resolution of asset pricing puzzles, *Journal of Finance* 59, 1481-1509.

Basak, Suleyman, and Domenico Cuoco, 1998, An equilibrium model with restricted stock market participation, *Review of Financial Studies* 11, 309-341.

Baxter, Marianne, and Urban Jermann, 1997, The international diversification puzzle is worse than you think, *American Economic Review* 87, 170-180.

Berk, Jonathan, 2000, Sorting out sorts, *Journal of Finance* 55, 407-427.

Black, Fisher, 1974, International capital market equilibrium with investment barriers, *Journal of Financial Economics* 1, 337-352.

Boldrin, Michelle, Lawrence Christiano, and Jonas Fisher, 2001, Habit persistence, asset returns, and the business cycle, *American Economic Review* 91, 149-166.

Brennan, Michael, Tarun Chordia, and Avanidhar Subrahmanyam, 1998, Alternative factor specifications, security characteristics, and the cross-section of expected stock returns, *Journal of Financial Economics* 49, Issue 3, 345-373.

Burmeister, Edwin, and Marjorie McElroy, 1988, Joint estimation of factor sensitivities and risk premia for the Arbitrage Pricing Theory, *Journal of Finance* 43, 721-735.

Campbell, John, 2000, Asset Pricing at the Millennium, *Journal of Finance* 55, 1515-1567.

Campbell, John, and John Cochrane, 1999, By force of habit: A consumption based explanation of aggregate stock market behavior, *Journal of Political Economy* 107, 205-251.

Chan, K.C., Andrew Karolyi, and Rene Stulz, 1992, Global financial markets and the risk premium on US equity, *Journal of Financial Economics* 32, 137-167.

Chan, Yeung, and Leonid Kogan, 2002, Catching up with the Joneses: Heterogeneous preferences and the dynamics of asset prices, *Journal of Political Economy* 110, 1255-1285.

Chen, Xiahong, and Sydney Ludvigson, 2005, Land of addicts? An empirical investigation of habit-based asset pricing models, forthcoming in the *Journal of Applied Econometrics.*

Cho, Chinhyung, Cheol Eun, and Lemma Senbet, 1986, International arbitrage pricing theory: An empirical investigation, *Journal of Finance* 41, 313-330.

Cocco, Joao, Francisco Gomes, and Pascal Maenhout, 2005, Consumption and portfolio choice over the life-cycle, *Review of Financial Studies* 18, 491-533.

Connor, Gregory, and Robert Korajczyk, 1993, A test for the number of factors in an approximate factor model, *Journal of Finance* 48, 1263-1291.

Daniel, Kent, and Sheridan Titman, 2005, Testing factor-model explanations of market anomalies, Working Paper, Northwestern University.

DeMarzo, Peter, Ron Kaniel, and Ilan Kremer, 2004, Diversification as a public good: Community effects in portfolio choice, *Journal of Finance* 59, 1677-1716.

DeMarzo, Peter, Ron Kaniel, and Ilan Kremer, 2008, Relative Wealth Concerns and Financial Bubbles, *Review of Financial Studies* 21, 19 - 50.

Dumas, Bernard, Campbell Harvey, and Pierre Ruiz, 2003, Are correlations of stock returns justified by subsequent changes in national output?, *Journal of International Money and Finance* 22, 777-811.

Elton, Edwin, Martin Gruber, and Christopher Blake, 1995, Fundamental economic variables, expected returns, and bond fund performance, *Journal of Finance* 50, 1229-1256.

Epstein, Larry, and Stanley Zin, 1989, Substitution, risk aversion and the temporal behavior of consumption and asset returns: A theoretical framework, *Econometrica*

57, 937-969.

Errunza, Vihang, and Etienne Losq, 1985, International asset pricing under mild segmentation: Theory and tests, *Journal of Finance* 40, 105-124.

Fama, Eugene, and John MacBeth, 1973, Risk, return and equilibrium: empirical tests, *Journal of Political Economy* 81, 607-636.

Fama, Eugene, and Kenneth French, 1998, Value versus growth: The international evidence, *Journal of Finance* 53, 1975-1999.

Ferson, Wayne, and George Constantinides, 1991, Habit persistence and durability in aggregate consumption: empirical tests, *Journal of Financial Economics* 29, 199-240.

Ferson, Wayne, and Campbell Harvey, 1994, Sources of risk and expected returns in global equity markets, *Journal of Banking and Finance* 18, 775-803.

Galí, Jordi, 1994, Keeping up with the Joneses: Consumption externalities, portfolio choice, and asset prices, *Journal of Money, Credit and Banking* 26, 1-8.

Gollin, Douglas, 2002, Getting Income Shares Right, *Journal of Political Economy* 110, 458-474.

Gomes, Francisco, and Alexander Michaelides, 2008, Asset pricing with limited risk sharing and heterogeneous agents, *Review of Financial Studies* 21, 415 - 448.

Gómez, Juan-Pedro, 2007, The impact of keeping up with the Joneses behavior on asset prices and portfolio choice, *Finance Research Letters* 4, Issue 2, 95-103.

Greene, William, 2007, *Econometric Analysis* (Prentice-Hall, Upper Saddle River, NJ).

Griffin John, 2002, Are the Fama and French factors global or country specific? *Review of Financial Studies* 15, 783-804.

Grilli, Vittorio, and Graciela Kaminsky, 1991, Nominal exchange rate regime and the real exchange rate: Evidence from the United States and Great Britain, 1885-1986, *Journal of Monetary Economics* 27, 191-212.

Gultekin, Mustafa, Bulent Gultekin, and Alessandro Penati, 1989, Capital Controls and International Capital Market Segmentation: The Evidence from the Japanese and American Stock Markets, *Journal of Finance* 44, 849-869.

Haliassos, Michael, and Alexander Michaeledis, 2002, Calibration and computation of household portfolio models, in Luigi Guiso, Michael Haliassos, and Tullio Jappelli, eds.: *Household Portfolios* (MIT Press).

Harvey, Campbell, 1991, The world price of covariance risk, *Journal of Finance* 46, 111-158.

Heaton, John, and Deborah Lucas, 2000, Portfolio choice and asset prices: the importance of entrepreneurial risk, *Journal of Finance* 55, 1163-1198.

Head, Allen, and Gregor Smith, 2003, The CCAPM meets Euro-interest rate persistence, 1960-2000, *Journal of International Economics* 59, 349-366.

Jagannathan, Ravi, and Zhenyu Wang, 1996, The conditional CAPM and the cross-section of expected returns, *Journal of Finance* 51, 3-53.

King, Mervyn, Enrique Sentana, and Sushil Wadhwani, 1994, Volatility and links between national stock markets, *Econometrica* 62, 901-933.

Korajczyk, Robert, and Claude Viallet, 1989, An empirical investigation of international asset pricing, *Review of Financial Studies* 2, 553-586.

Lauterbach, Beni, and Haim Reisman, 2004, Keeping up with the Joneses and the home bias, *European Financial Management* 10, 225-234.

Lettau, Martin, and Sydney Ludvigson, 2001, Resurrecting the (C)CAPM: A cross-sectional test when risk premia are time-varying, *Journal of Political Economy* 109, 1238-1287.

Lewellen, Jonathan, Stefan Nagel, and Jay Shanken, 2006, A skeptical appraisal of asset-pricing tests, working paper, Dartmouth University.

Lo, Andrew, and Craig MacKinlay, 1990, When are contrarian profits due to stock market overreaction? *Review of Financial Studies* 3, 175-205.

Papell, David, 1997, Searching for stationarity: Purchasing power parity under the current float, *Journal of International Economics* 43, 313-332.

Polkovnichenko, Valery, 2004, Limited stock market participation and the equity premium, *Finance Research Letters* 1, 24-34.

Priestley, Richard, 1997, Seasonality, stock returns and the macroeconomy, *Economic Journal* 107, 1742-1750.

Ravina, Enrichetta, 2005, Keeping up with the Joneses: evidence from micro data, working paper, Columbia Business school.

Roll, Richard, 1977, A critique of the asset pricing theory's tests, *Journal of Financial Economics* 4, 129-176.

Shanken, Jay, and Guofu Zhou, 2006, Small-sample properties of estimators and tests for Beta-pricing models: A simulation, Emory University, Working Paper.

Shore, Stephen, and Joshua White, 2002, External habit formation and home bias puzzle, University of Pennsylvania, Working Paper.

Solnik, Bruno, 1974, An equilibrium model of the international capital market, *Journal of Economic Theory* 8, 500-524.

Stulz, René v , 1981, A model of international asset pricing, *Journal of Financial Economics* 9, 383-406.

Viceira, Luis, 2001, Optimal portfolio choice for long-horizon investors with non-tradable labor income, *Journal of Finance* 56, 433-470.

Weil, Philippe, 1989, The equity premium puzzle and the risk-free rate puzzle, *Journal of Monetary Economics* 24, 401-421.

Wu, Yangru, 1996, Are real exchange rates nonstationary? Evidence from a panel-data test, *Journal of Money, Credit, and Banking* 28, 54-63.

Endnotes

[1]Burmeister and McElroy (1988) and Elton, Gruber, and Blake (1995) both use joint estimation methods to test asset pricing models. Lewellen, Nagel, and Shanken (2006) suggest that the use of individual securities in tests of asset pricing models would eliminate the problems caused by portfolio formation based on, for example, size and book-to-market.

[2]For the exogenous Joneses model we need to assume a common risk aversion coefficient across countries.

[3]The price of nondiversifiable wealth endowments is determined by market clearing in the corresponding spot markets.

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

[5]Notice that by rewriting $\gamma = (\alpha - 1)/\alpha$, $\alpha \geq 1$, the utility function (1) becomes the standard "ratio-habit" representation in Abel (1990):

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

with average consumption elasticity of marginal utility $(\alpha - 1)$. Our formulation is more convenient for the empirical tests of the model.

[6]Note that $u_c(c, w/c) \approx u_c(c^0, w^0/c^0) + u_{c,c}(c^0, w^0/c^0)(c - c^0) + u_{c,w/c}(c^0, w^0/c^0)(w/c - w^0/c^0) = u_c(1, w^0)\left[1 - \alpha(1-\gamma)(r'x + R - 1) + \frac{w^0}{1+w^0}\alpha\gamma(r'X^w + R - 1)\right]$. An approximation to the first-order condition is obtained by multiplying this expression by r and

taking expectations. The linear approximation for the optimal portfolio follows after solving for x^* and taking into account that $E(r'r) \approx \Omega$ and -$E(r)(R-1) \approx 0$ for small values of $E(r)$ and the (net) risk-free rate, $R-1$.

[7]We use the single-parameter, CRRA utility function. It can be shown that the model yields the same predictions for the cross-section of stock returns if this utility function is replaced by the two-parameter, CES utility function in Section IV.B of DeMarzo, Kaniel, and Kremer (2004), page 1699. It is clear that the exogenous specification has an advantage in this comparison, since it has two parameters and therefore an additional degree of freedom.

[8]Note that, like in the exogenous specification, we can approximate $u_c(c,p)$ around the initial endowment/price (c^0, p^0) such that

$$u_c(c,p) \approx u_c(1, \delta \bar{w}^{-\alpha}) \left[1 - \alpha(r'x + R - 1) + \frac{\bar{w}^0}{1+\bar{w}^0}(\alpha - 1)\left(\frac{\bar{w}p}{\bar{w}^0} - 1\right) \right].$$

We replace $\frac{\bar{w}p}{\bar{w}^0} = R + r'X^w$ in the equation. After multiplying by r and taking expectations we obtain the investor's first-order condition for the optimal portfolio. The same approximations as in the exogenous specification are assumed to linearize the solution.

[9]Galí (1994) has shown that if all agents in both countries hold well-diversified portfolios, that is, if there is no constraint ($\theta_k = 0$, $k = \{l, f\}$), the resulting equilibrium is symmetric with no keeping up with the Joneses behavior.

[10]The Appendix shows the effect of the orthogonalization and compares the prices of risk and the betas of our estimation vis-à-vis the regression without orthogonalizing the

return on the local, non-diversifiable income.

[11]In a simplified setting, De Marzo, Kaniel, and Kremer (2004) show that the endogenous Joneses equilibrium can be multiple. If there exists a large enough fraction of investors who are exogenously biased the equilibrium becomes unique. The case discussed here can be interpreted as a limit case where *all investors* are exogenously biased (i.e., exhibit keeping up with the Joneses preferences). Hence, the exogenous Joneses equilibrium is unique.

[12]We are limited in the number of portfolios that we can form using the double sorting procedure because of the relatively small number of stocks at the beginning of the sample. By forming nine portfolios we ensure that there are a minimum of 16 stocks in each portfolio in Germany, the country with the fewest listed stocks. This should provide a well-diversified portfolio return.

[13]Connor and Korajczyk (1993) argue that residuals may be cross-correlated due to industry-specific factors that are not pervasive across the whole cross-section.

[14]Given $\omega_k \theta_k b_k$ in (14), the coefficient of interest, b_k, cannot be independently estimated without additional moment conditions on the distribution of financial wealth across countries ω_k, and more importantly, the ratio of nondiversifiable to total country wealth, θ_k.

[15]We thank the referee for this suggestion. As explained in footnote 5, assuming $\gamma = \frac{\alpha-1}{\alpha}$ results in the "ratio-habit" representation in Able (1990). This would reduce the Joneses inferences to a single parameter, α, just like in the endogenous model. Notice, however, that under this specification, $b = \alpha - 1$ and $\tau = 1$. This implies $H^{-1} =$

$\sum_k \omega_k \tau_k = 1$. H is estimated endogenously; see equation (14). Assuming the ratio-habit representation is equivalent to introducing an ad hoc restriction on H inconsistent with the estimation procedure. When we use individual securities, where we obtain a reasonable estimate of the market price of risk, the inferred value of $H = 2.1672$ is clearly different from one. This implicitly rejects the ratio-habit formulation.

[16]This was to be expected since, by construction, the country-specific risk factors are orthogonal to the market portfolio return.

Table I
Fama and MacBeth Cross-sectional Tests

The table reports Fama and MacBeth (1973) second stage cross-sectional estimates of prices of risk (%) using 10 size portfolios from each of the U.S., U.K., Japan (JP), and Germany (GM) and nine size and lagged factor loading portfolios for each country. λ^M is the world stock market price of risk, λ^{US} is the orthogonal U.S. local labor income price of risk, λ^{UK} is the orthogonal U.K. local labor income price of risk, λ^{JP} is the orthogonal Japanese local labor income price of risk, λ^{GM} is the orthogonal German local labor income price of risk, λ^{hml} is the international version of Fama and French's high minus low book-to-market factor price of risk, λ^{cb} is the currency basket price of risk, λ^{i} is the international inflation price of risk, λ^{ip} is the international industrial production price of risk, and λ^{m} is the international interest rate price of risk. R^2 is $\left[Var_c(\overline{r}_i) - Var_c(\overline{e}_i)\right]/Var_c(\overline{r}_i)$ and $\overline{R}^2$ is the adjusted R^2. t-statistics are reported in parentheses.

Model	λ^M	λ^{US}	λ^{UK}	λ^{JP}	λ^{GM}	λ^{hml}	λ^{cb}	λ^{i}	λ^{ip}	λ^{m}	$\overline{R}^2$
1	−0.397	−0.052	−0.218	−0.148	−0.537						31
	(1.13)	(0.80)	(2.02)	(0.73)	(2.74)						
2	−0.326										4
	(0.79)										
3	−0.369					0.352					20
	(0.91)					(1.33)					
4	−0.531						−0.421				21
	(1.30)						(1.86)				
5							−0.450	0.008	−0.030	−0.071	21
							(2.09)	(0.32)	(0.18)	(0.28)	

Table II
NLSUR Estimates Using Individual Stock Returns

The table reports NLSUR estimates of prices of risk (%) using individual stock returns. The lambdas and $\overline{R}^2$ are estimated as in Table I. t-statistics are reported in parentheses. The table reports the mean, standard deviation (Std Dev), 25th percentile and 75th percentile from one hundred estimates of the parameters of interest. KEEPM is the Keeping-up-with-the-Joneses model, ICAPM is the international CAPM, IFF is the international Fama and French model, ICAPM_CR is the ICAPM that includes a currency risk factor, IMACRO is an international multi-factor model.

Panel A: KEEPM

	λ^M	λ^{US}	λ^{UK}	λ^{JP}	λ^{GM}	λ^{hml}	λ^{cb}	λ^i	λ^{ip}	λ^m	$\overline{R}^2$
Mean	0.574	−0.598	−1.007	−0.519	−0.531						51
	(4.65)	(6.03)	(6.63)	(0.81)	(1.94)						
Std Dev	0.266	0.552	1.035	2.098	1.033						24
	(1.74)	(3.13)	(2.16)	(2.34)	(1.98)						
25th %	0.528	−0.753	−1.067	−0.717	−0.727						30
	(3.82)	(8.22)	(8.17)	(3.12)	(3.35)						
75th %	0.696	−0.362	−0.677	0.265	−0.091						71
	(5.86)	(3.81)	(5.23)	(1.19)	(0.51)						

Panel B: ICAPM

	λ^M	λ^{US}	λ^{UK}	λ^{JP}	λ^{GM}	λ^{hml}	λ^{cb}	λ^i	λ^{ip}	λ^m	$\overline{R}^2$
Mean	0.585										< 0
	(6.63)										
Std Dev	0.069										15
	(0.84)										
25th %	0.542										< 0
	(5.76)										
75th %	0.634										6
	(6.95)										

Panel C: IFF

	λ^M	λ^{US}	λ^{UK}	λ^{JP}	λ^{GM}	λ^{hml}	λ^{cb}	λ^i	λ^{ip}	λ^m	$\overline{R}^2$
Mean	0.665					−0.863					3
	(6.18)					(4.24)					
Std Dev	0.119					0.258					17
	(1.03)					(1.12)					
25th %	0.576					-1.010					< 0
	(5.45)					(5.16)					
75th %	0.731					−0.680					12
	(6.94)					(3.46)					

Panel D: ICAPM_CR											
	λ^M	λ^{US}	λ^{UK}	λ^{JP}	λ^{GM}	λ^{hml}	λ^{cb}	λ^i	λ^{ip}	λ^m	$\overline{R}^2$
Mean	0.589 (5.98)						−0.417 (0.98)				12
Median	0.585 (6.07)						−0.185 (0.95)				12
Std Dev	0.107 (1.08)						1.327 (2.48)				24
25th %	0.545 (5.31)						−0.721 (2.91)				< 0
75th %	0.644 (6.66)						0.558 (1.27)				28

Panel E: IMACRO											
	λ^M	λ^{US}	λ^{UK}	λ^{JP}	λ^{GM}	λ^{hml}	λ^{cb}	λ^i	λ^{ip}	λ^m	$\overline{R}^2$
Mean							0.536 (1.18)	−0.209 (1.74)	−0.690 (1.43)	−0.661 (1.18)	69
Std Dev							2.995 (24.67)	0.630 (17.91)	2.568 (21.49)	3.083 (21.33)	20
25th %							−0.783 (1.40)	−0.185 (3.21)	−1.342 (3.19)	−2.245 (2.89)	60
75th %							2.147 (3.02)	−0.003 (0.67)	0.007 (0.37)	0.384 (0.65)	80

Table III
Pricing Errors Using Portfolio Returns

This table reports the analysis of pricing errors (%) using portfolio returns. KEEPM is the Keeping-up-with-the-Joneses model, ICAPM is the international CAPM, IFF is the international Fama and French (1998) model, ICAPM_CR is the ICAPM that includes a currency risk factor, and IMACRO is an international multifactor model. S denotes 10 size sorted portfolios and SFL denotes nine size and factor loadings sorted portfolios. All includes all 76 portfolios. Panel A reports the square root of the average squared pricing error for the aggregated portfolios. Standard errors are reported in parentheses. Panel B reports a Wald test that the pricing errors of the portfolios that make up the aggregate portfolios are jointly zero. The final row of Panel B reports a test that the pricing errors from the KEEPM model are smaller than those from the other models. * indicates that the null hypothesis is rejected at the 5% level and ** indicates that the null hypothesis is rejected at the 10% level.

Panel A: Square Root of Average Squared Pricing Errors					
Portfolios	KEEPM	ICAPM	IFF	ICAPM_CR	IMACRO
US S	0.188	0.203	0.160	0.151	0.197
	(0.19)	(0.21)	(0.16)	(0.25)	(0.28)
US SFL	0.440	0.627	0.561	0.490	0.420
	(0.39)	(0.42)	(0.39)	(0.39)	(0.34)
UK S	0.460	0.631	0.554	0.480	0.353
	(0.33)	(0.38)	(0.43)	(0.33)	(0.45)
UK SFL	0.148	0.140	0.150	0.151	0.230
	(0.11)	(0.07)	(0.11)	(0.09)	(0.16)
JP S	0.298	0.376	0.323	0.433	0.430
	(0.33)	(0.30)	(0.29)	(0.29)	(0.32)
JP SFL	0.291	0.231	0.241	0.220	0.263
	(0.25)	(0.17)	(0.161)	(0.19)	(0.21)
GM S	0.177	0.132	0.155	0.149	0.116
	(0.15)	(0.16)	(0.16)	(0.11)	(0.08)
GM SFL	0.349	0.410	0.332	0.457	0.421
	(0.16)	(0.16)	(0.16)	(0.22)	(0.18)
All	0.290	0.343	0.310	0.320	0.302
	(0.27)	(0.32)	(0.25)	(0.25)	(0.23)
Panel B: Tests of Pricing Errors					
Portfolios	KEEPM	ICAPM	IFF	ICAPM_CR	IMACRO
US S	68.115*	63.367*	53.614*	72.796*	83.434*
US SFL	207.14*	195.06*	169.96*	177.97*	183.67*
UK S	93.405*	75.206*	90.874*	74.133*	91.775*
UK SFL	17.975*	18.422*	23.023*	16.743**	28.237*
JP S	50.875*	44.502*	42.938*	46.386*	75.053*
JP SFL	69.156*	29.600*	28.388*	36.468*	45.376*
GM S	16.961**	12.820	15.802**	17.008**	10.455
GM SFL	45.081*	19.614*	18.465*	36.491*	33.947*
All	2898.01*	2553.15*	2406.45*	2624.43*	2606.81*
All Keep - other		−3.161*	−1.015	−1.457	-0.533

Table IV
Pricing Errors Using Individual Securities

This table reports analysis of pricing errors (%) using individual securities. KEEPM is the Keeping-up-with-the-Joneses model, ICAPM is the international CAPM, IFF is the international Fama and French (1998) model, ICAPM_CR is the ICAPM that includes a currency risk factor, and IMACRO is an international multifactor model. All includes the average pricing errors across 80 individual stocks and across 100 estimates of the model. Panel A reports the square root of the average squared pricing error for aggregated portfolios. Standard errors are reported in parentheses. Panel B reports a *t*-test that the average pricing error over the one hundred versions of the model are jointly zero. The final row of Panel B tests whether the pricing errors from the KEEPM model are smaller than those from the other models. * indicates that the null hypothesis is rejected at the 5% level and ** indicates that the null hypothesis is rejected at the 10% level.

Panel A: Square Root of Average Squared Pricing Errors					
Models	KEEPM	ICAPM	IFF	ICAPM_CR	IMACRO
ALL	0.236	0.363	0.353	0.332	0.210
	(0.06)	(0.03)	(0.03)	(0.04)	(0.33)
Panel B: Tests of Pricing Errors					
Models	KEEPM	ICAPM	IFF	ICAPM_CR	IMACRO
ALL	36.111*	119.99*	111.73*	75.290*	14.715*
All KEEP - Other		-19.406*	-17.877*	-14.666*	3.993*

Table V
Inference on the Model Parameters Using Portfolios

The table reports the inferred deeper parameters in the model using the prices of risk estimated with portfolios. We consider in each panel a different price of risk for the global market portfolio, λ_M (%). We then estimate the prices of risk for the orthogonal risk factors. The implied aggregate risk aversion coefficient, H, is also reported. θ is the proportion of labor income over the total country income. The two values per country are taken from Table II on pg. 43 in Gollin (2002). The higher value includes the compensation of the self-employed. The parameter b represents a measure of portfolio bias. It is a function of the Joneses parameter, γ, in the exogenous model; in the endogenous model, it is a function of the relative risk aversion coefficient for the representative agent, α_{EN}. α_{EX} is the implied risk aversion coefficient in the exogenous model without Joneses behavior. $\tau^{-1} = \alpha_{EX}(1-\gamma)$ is the scaled risk aversion coefficient for the representative agent in the presence of Joneses behavior.

θ	b	γ	α_{EN}	α_{EX}	τ^{-1}
Panel A: $\lambda_M = 0.215$; $H = 0.6683$					
United States					
60%	20.71	0.9539	-0.0507	100.67	4.63
77%	16.14	0.9416	-0.0660	72.15	4.20
United Kingdom					
57%	391.61	0.9974	-0.0025	100.67	0.25
81%	275.57	0.9963	-0.0036	72.15	0.26
Japan					
56%	61.83	0.9840	-0.0164	100.67	1.60
73%	47.43	0.9793	-0.0215	72.15	1.48
Germany					
57%	1370.12	0.9992	-0.0007	100.67	0.073
81%	964.16	0.9989	-0.0010	72.15	0.074
Panel B: $\lambda_M = 0.43$; $H = 1.9640$					
United States					
60%	7.20	0.8781	-0.1610	100.80	12.28
77%	5.61	0.8488	-0.2166	72.72	10.99
United Kingdom					
57%	128.62	0.9922	-0.0078	100.80	0.77
81%	90.51	0.9890	-0.0111	72.72	0.79
Japan					
56%	24.86	0.9613	-0.0419	100.80	3.89
73%	19.07	0.9501	-0.0553	72.72	3.62
Germany					
57%	452.75	0.9977	-0.0022	100.80	0.222
81%	318.60	0.9968	-0.0031	72.72	0.227
Panel C: $\lambda_M = 0.645$; $H = 3.2299$					
United States					
60%	4.47	0.8174	-0.2875	100.92	18.42
77%	3.48	0.7772	-0.4017	73.28	16.32
United Kingdom					
57%	75.46	0.9869	-0.0134	100.92	1.31
81%	53.10	0.9815	-0.0191	73.28	1.35
Japan					
56%	17.39	0.9456	-0.0610	100.92	5.48
73%	13.34	0.9302	-0.0810	73.28	5.11
Germany					
57%	267.32	0.9962	-0.0037	100.92	0.37
81%	188.11	0.9947	-0.0053	73.28	0.38

Table VI
Inference on the Model Parameters Using Individual Securities

The table reports the inferred deeper parameters in the model using the mean prices of risk estimated with random samples of individual securities. Panel A uses the mean prices of risk as reported in Table II, Panel A. The implied aggregate risk aversion coefficient is $H = 2.1672$. For comparison purposes, Panel B reports the parameter estimates after imposing $\lambda_{M} = 0.43\%$ as in Table V, Panel B. The implied aggregate risk aversion coefficient is $H = 1.3135$. Parameters $\theta, b, \gamma, \theta, \alpha_{EN}, \alpha_{EX}$, and τ^{-1} are as in Table V.

θ	b	γ	α_{EN}	α_{EX}	τ^{-1}
Panel A: $\lambda_M = 0.574$; $H = 2.1672$					
United States					
60%	174.13	0.9942	-0.0057	510.92	2.91
77%	135.69	0.9926	-0.0074	379.00	2.77
United Kingdom					
57%	812.89	0.9987	-0.0012	510.92	0.62
81%	572.03	0.9982	-0.0017	379.00	0.66
Japan					
56%	19.09	0.9502	-0.0552	510.92	25.42
73%	14.64	0.9360	-0.0732	379.00	24.21
Germany					
57%	217.20	0.9954	-0.0046	510.92	2.34
81%	152.84	0.9935	-0.0065	379.00	2.46
Panel B: $\lambda_M = 0.43$; $H = 1.3135$					
United States					
60%	287.74	0.9965	-0.0034	510.30	1.76
77%	224.21	0.9955	-0.0044	378.33	1.67
United Kingdom					
57%	1340.78	0.9992	-0.0007	510.30	0.38
81%	943.51	0.9989	-0.0010	378.33	0.40
Japan					
56%	30.93	0.9686	-0.0334	510.30	15.98
73%	23.72	0.9595	-0.0439	378.33	15.29
Germany					
57%	359.25	0.9972	-0.0027	510.30	1.41
81%	252.81	0.9960	-0.0039	378.33	1.49