Characterizing Asymmetric Information in International Equity Markets

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Abstract

This paper studies international portfolio flows of US investors to examine the information structure of international equity markets. Based on a model of portfolio choice with both public and private information, we propose new empirical measures of trades due to private information. We show that these trades help explain the cross section of international equity returns, after controlling for public information. We find that such trades are highly correlated across countries. In particular, a common "global" factor accounts for about half of the variation in trades due to private information and can also be used to predict returns in many countries. The finding that a substantial portion of trades due to private information across countries contains the same common information challenges the conventional view that domestic investors have better private information about their home market than foreign investors.

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1 Introduction

Recently, many authors have argued that asymmetric information is a key determinant of international equity flows and returns.[1] This view is based on two stylized facts. The first is that the level of international portfolio investment is well below that suggested by models of international diversification (the "home bias" puzzle).[2] The second is that foreign investors tend to hold those domestic securities that are more familiar to them.[3] Beyond these facts, not much is known about the information structure in international capital markets. Nevertheless, the conventional view is that domestic investors have better information about their home market than do foreign investors. Private information is typically taken to mean knowledge of future firm-specific payoffs or country-specific information about economic or political conditions.[4]

In this paper, we provide new evidence about the information structure in international equity markets by comparing trades due to private information across the major developed countries. Based on a model of portfolio choice, we provide new measures of private information in US investors' international equity flows. Under the conventional view in the literature, we would not expect private information about one country to be useful for forecasting returns in other countries. Thus, US investors' trades due to private information should not be highly correlated across countries. We find that, contrary to the conventional view, trades due to private information are highly correlated across countries. In particular, we identify a common "global" factor which accounts for approximately half of the variation in trades due to private information.[5] This common factor is useful for forecasting returns in many countries and has an economically large impact on realized return variation. To the best of our knowledge, this is the first time that common private information has been shown to exist in international markets.

One interpretation of our common factor in private information is that it captures US investors' superior information about the US economy or US asset markets. This is what we would expect under the conventional view that domestic investors have superior

[1] For a survey on international capital flows and returns see Stulz (1999).

[2] See, French and Poterba (1991), Gehrig (1993), Shukla and van Inwegen (1995), Brennan and Cao (1997) and Coval (1999)) for models of home bias with asymmetric information. For a survey of the home bias literature, see Lewis (1999).

[3] See Kang and Stulz (1997), Choe, Kho, and Stulz (1999), Dahlquist and Robertson (2001), and Ahearne, Griever, and Warnock (2001). We return to this evidence below.

[4] Note that this definition of private information is broader than "inside information" about a particular firm. In the international literature, private information may concern payoffs on individual firms but is more likely concerned with country-wide phenomena. We suggest possible sources of this below.

[5] This global factor in private information is obtained after removing the effects of global factors in public information as detailed below.

information about their home market. Our point is that this information is likely to have value in international markets as well. The sizeable volume of international trade and foreign direct investment originated in the US makes the US one of the largest trading partners for countries around the globe. In addition, papers testing international asset pricing models show that investors' public information can be captured by a small set of global factors (predominately US instruments) which drive US and foreign asset returns.[6] It is thus not implausible that private information signals on US equities also have information content for foreign equities.[7] We consider some possible economic explanations in Section 4.3 below.

According to our results, US investors are likely to play a special role in international markets. Rather than suffering from an information disadvantage because of insufficient local information, they may enjoy an information advantage because of superior global information. At first sight, this view appears inconsistent with the home-bias puzzle. However, it is possible that being relatively better informed about their home market causes US investors to tilt their portfolios towards that market, or to securities more closely related to it.[8] In other words, the 'comparative information advantage' which makes them hold the majority of US equities is consistent with an absolute advantage in holding foreign equities.

Our estimates are constructed from data on US investors' equity flows in the eight largest foreign equity markets. Changes in the share of a foreign market held by US investors must be due to either public or private signals. The literature that examines international equity returns has uncovered many public information variables that are useful in explaining the cross section of expected returns.[9] We use these variables to extract the trades of US investors based on public information. The residuals from these regressions represent (noisy) measures of trades due to US investors' private information. We label these initial estimates our "broad" measures of private information.

Portfolio theory predicts that trades due to private information must forecast returns, after controlling for public information. This leads to a test of our measure. We use the latent-factor model of Hansen and Hodrick (1980) and Gibbons and Ferson (1985) to model the cross section of expected international equity returns based on public informa-

[6] See, for example, Harvey (1991), Ferson and Harvey (1993), and Campbell and Hamao (1992). This view is contentious, however; see Griffin (2001).

[7] The domestic asset pricing literature has also illustrated the potential role for common factors in private information to explain some time series properties of returns (e.g. Subrahmanyam (1991) and Chan 1993).

[8] The evidence cited below that shows that foreign (i.e. US) investors tilt their foreign asset holdings towards firms that are larger or have a greater presence in international markets is thus also consistent with the new view.

[9] See Karolyi and Stulz (2001) for a survey.

tion. We then show that an unexpected net flow by US investors into a foreign market has a statistically significant positive effect on the market's return in the current and following three months, holding the public information effects constant. The private information effects are large economically as they can account for a significant amount of the *realized* return variation in the countries that we examine.

To obtain a long time span of US investor portfolio flows into the developed countries, we use monthly flow data from the US Treasury, which we detail below. Using monthly data to extract private information contains a potentially serious identification problem: our broad measure may overstate the role of private information as trades due to (public) news released during the month are counted as trades "due to private information".[10] To solve this problem, we construct an alternative, "conservative" measure of private information by including contemporaneous (end-of-month) values of the public information variables and *returns* in the flow regressions. By construction, this measure is orthogonal to *any* public news released during the month that affects returns. Of course, this measure will understate the role of private information as all trades based on private information that increase prices during the month are now counted as trades due to public information. Nevertheless, under our conservative measure, which is by construction orthogonal to contemporaneous returns, an unexpected net flow by US investors into a foreign market has a statistically significant effect on the market's return in the following month. These effects are also large economically according to our realized return variation metric.

Relative to the literature, the paper makes three main contributions. The first is to provide a simple model of (aggregate) portfolio positions made by a set of investors with both private and public information. Here we build on the setups of Admati (1985), Gehrig (1993), Brennan and Cao (1997), and Coval (1999). We use the model to justify our regression approach and clarify the various biases that can occur. We also show that our measures of private information should appear as an additional factor in a pricing model, together with factors estimated from public information.

The second contribution is to construct empirical measures of trades due to private information. A limitation of the existing empirical literature is the use of data sets that are specific to narrow groups of investors or time periods. Here, we extract the private information from a comprehensive data set of monthly purchases and sales of equities in the eight major foreign markets by US investors over the 1977 to 2000 period. The low-frequency nature of our data requires that we extend the market microstructure work of Hasbrouck (1991a,b) to allow for news about public information to affect both equity

[10] For example, US investors could be "trend followers" purchasing foreign equities in response to increasing foreign equity returns during the month (Froot et al. 2001).

flows and returns. We find that the standard instruments used to predict international equity returns also predict *gross* equity flows (purchases and sales) by US investors in these markets. We use our model of gross flows to construct a measure of unanticipated net flows, which we label the private information of US investors. We provide evidence to show that our method of forecasting net flows from separate forecasts of gross purchases and sales results in a better measure of expected net flows than by forecasting them directly. Our modeling of expected net flows due to public information improves on the approaches in Bohn and Tesar (1996a,b), Brennan and Cao (1997), and Portes and Rey (2000).

The third contribution is to examine the factor structure of the private information. Similar to the analyses of expected returns based on public information, we are interested in describing the extent to which measured private information displays significant global components. The model indicates that a simple factor analysis can reveal these components and that tests of their impacts on the cross section of returns can be performed. We emphasize that such a factor analysis requires a comprehensive data set like the one used in this paper. Our empirical tests show that a common or global factor exists in US investors' private information. This common factor accounts for roughly 55 per cent of the total variation in private information of US investors. Shocks to the common factor produce persistent price impacts in the eight foreign markets analyzed as well as in the US equity market. This latter result supports the view that the common factor is derived from the signal that US investors get about US asset and goods markets. These shocks to global component of private information are also economically large. Our results thus confirm that private information matters, but they challenge the conventional (i.e. country-specific) view of why it matters.

The paper is related to a growing literature which examines the effect of private information on portfolio stocks and flows. A number of papers have shown that foreign investors tend to hold those domestic securities that are more familiar to them. For example, Kang and Stulz (1997) find that foreign investors in Japan tend to hold shares of larger firms and firms with significant export sales. They argue that investors are likely to have more information about these firms (see also Choe, Kho, and Stulz (1999) for Korea). Similarly, Dahlquist and Robertson (2001) examine foreign ownership of Swedish firms and show that foreigners display a preference for firms with a significant presence in international markets (firms with a foreign listing or with high export sales), controlling for firm size. Using a sample of 48 countries, Ahearne, Griever, and Warnock (2001) find that the share of a country's stock market that is publicly listed in the United States (a measure inversely related to private information) is an important determinant of the bias in US foreign stock ownership, even after controlling for a variety of instruments including transactions costs of trading in the country and direct barriers to foreign ownership.[11] As discussed earlier, these results about foreign equity holdings are not inconsistent with our results about foreign equity flows.

Another set of papers has examined whether foreign investors underperform domestic investors. On the one hand, Frankel and Schmukler (1996) have shown that, prior to the December 1994 Mexican crisis, the net asset values of closed-end funds dropped faster than the prices of the funds suggesting that Mexican investors had better knowledge of the likelihood of the crisis. On the other hand, Froot and Ramadorai (2001) show that unexpected inflows into closed-end funds cause an increase in the prices of both the net asset value of the fund and that of the fund itself indicating that foreign investors have significant private information. Choe, Kho, and Stulz (2001) analyze the trading behavior of foreign investors (US and others) and domestic institutions and individuals around days of significant abnormal returns and days of large buying or selling activity in Korea. They find that foreign investors trade at worse prices relative to domestic individuals only. However, the private information advantage of domestic individuals is minimal relative to an investor that holds stocks for a substantial period of time.[12] Our findings suggest a reason for the good performance of US investors in some of these studies.

The paper proceeds as follows. In Section 2, we introduce a general model of portfolio choice under asymmetric information and outline our empirical approach. In Section 3, we present our data and our models of the expected components of international equity flows and returns. In Section 4, we detail how private information contains global components and show that our measures of private information can be used to forecast returns. We discuss how our results are robust to alternative interpretations and give some economic motivations for them. In Section 5, we conclude.

[11] See Coval and Moskowitz (1999) and Huberman (2001) for similar effects in domestic equity markets. There is also evidence that the volume of transactions between countries is correlated with measures of information flow (Portes and Rey (2000)).

[12] In addition, Seasholes (2000) finds that foreign investors in Taiwan systematically accumulate assets before positive earnings announcements and systematically sell assets before negative earnings announcements. Bailey and Mao (2001) analyze periods of earnings and dividends announcements in Thailand and Singapore and find evidence consistent with foreigners having superior information as compared to domestic residents. Hamao and Mei (2001) find no significant evidence that foreigners are able to time the Japanese stock market. However, Karolyi (1999) shows that foreign investors have outperformed domestic investors in Japan. Using Finnish data, Grinblatt and Keloharju (2000) foreign investors seem to outperform domestic household investors.

2 The Asymmetric Information Framework

In this section we outline our model and discuss how measures of private information can be obtained and empirical tests of private information effects may be performed.

2.1 The Portfolio Choice Model

There are N assets and a continuum of infinitesimally small investors. For simplicity each asset represents the stock index of a country. A typical investor i chooses streams of consumption c_t^i and asset holdings $\theta_t^i = (\theta_t^{i1}, ..., \theta_t^{iN})$ to maximize her expected lifetime utility

$$-E_0 \left[\sum_{t=0}^{\infty} \delta^t \exp^{-\rho c_t^i} \right],$$

subject to the sequence of period budget constraints

$$w_{t+1}^i = R^f \left(w_t^i - c_t^i \right) + \theta_t^{i\prime} R_{t+1} + y_{t+1}^i. \tag{1}$$

Here w_t^i is the beginning-of-period wealth, y_{t+1}^i represents non-asset income, and R_{t+1} is the vector of per-share excess returns on the risky assets (i.e., if P_t is the vector of ex-dividend share prices and D_t is the vector of dividends, then $R_{t+1} = P_{t+1} + D_{t+1} - R^f P_t$). All prices are measured in US dollars. The riskless interest rate is R^f, the discount rate is $\delta < 1$, and ρ is the coefficient of absolute risk aversion. All exogenous variables are assumed to be normally distributed. Finally, $E_0(.)$ is the conditional expectations operator.

Information Structure

Let $\mathcal{I}_t^i$ denote the investor's information set available at time t. It includes past and present realizations of asset holdings and non-asset income as well as two vectors of public information variables, Z_t and Y_t, which we distinguish below. In addition, $\mathcal{I}_t^i$ contains a vector of *private information signals*, σ_t^i. Signals are assumed to be iid across agents drawn from a normal distribution with time-varying first moments.

We assume that any investor's payoff-relevant information at time t can be summarized by a vector $I_t^i \in \mathcal{I}_t^i$ which evolves over time in a Markovian fashion:

$$I_{t+1}^i | I_t^i \sim N \left(\Phi_0 + \Phi_1 I_t^i, \Sigma_{II} \right). \tag{2}$$

We define payoff-relevant information as all the information that the investor would use in forecasting future values of asset returns and non-asset income. For concreteness, we write $I_t^i = (\sigma_t^{i\prime}, Y_t^\prime, Z_t^\prime, R_t^\prime, y_t^i, \hat{I}_t^\prime)^\prime$; i.e. I_t^i contains all current public and private signals. Here the vector $\hat{I}_t$ represents all information revealed by past signals that remains payoff relevant during the current month.[13] We assume that all such information is public. This implies that any private information diffuses quickly: no information can remain private for more than one period. The assumption that the matrices Φ_0, Φ_1, and Σ_{II} are independent of the investor's identity is natural given that we assume iid signals.

The information content of signals is implicit in the correlation matrix Σ_{II}. Signals can be directly related to returns or indirectly through other variables that are in turn correlated with returns (e.g. variables in $\hat{I}_t$). Also, each signal can be informative about a variety of assets. This can be the case if there is correlated private information as in the models of Subrahmanyam (1991) and Chan (1993).

Optimal Asset Allocation and Aggregation Across US Investors

It is straightforward to show that portfolio demand is linear in payoff relevant information:

$$\begin{aligned} \theta_t^i &= \overline{\theta} + \Theta I_t^i \\ &= \overline{\theta} + \Theta^z Z_t + \Theta^Y Y_t + \Theta^R R_t + \Theta^I \hat{I}_t + \Theta^y y_t^i + \Theta^\sigma \sigma_t^i, \end{aligned} \tag{3}$$

where the coefficient matrices are detailed in the Appendix. The net flow of investor i is given by $\theta_t^i - \theta_{t-1}^i = \Theta\left(I_t^i - I_{t-1}^i\right)$. In our data set, we observe aggregate net flows by *all* US investors, $\text{NF}_t^{US} \equiv \int_i (\theta_t^i - \theta_{t-1}^i) di$. Under the assumption that US investors are all identical,[14]

$$\text{NF}_t^{US} = \Theta^z \Delta Z_t + \Theta^Y \Delta Y_t + \Theta^R R_t + \Theta^I \Delta \hat{I}_t + \Theta^\sigma \Delta \sigma_t,$$

where $\sigma_t \equiv \int \sigma_t^i di$ is the average private signal. Here we have assumed that average non-traded income, $\int_i y_t^i di$ is contained in Z_t.

2.2 Measuring Trades due to Private Information

To obtain an empirical measure of aggregate private information from the net flows data we first have to specify the econometrician's information set and how it relates to the

[13] In general, it is not clear that a time-invariant "summary vector" exists in an equilibrium setup with private information. However, this is guaranteed if all past private information is revealed after finitely many periods, as in Townsend (1983). In Albuquerque et al. (2002) we exhibit a setup with this property in an equilibrium context.

[14] To solve this integral over a continuum of agents we proceed as in Admati (1985). Let $\tilde{\varepsilon}_i$, $i \in [0, 1]$ be almost surely integrable. Then, if ε_i is an independent random variable with zero mean and bounded variance $\int_i (\varepsilon_i + \tilde{\varepsilon}_i)\, di \equiv \int_i \tilde{\varepsilon}_i di$.

investor's information set. We assume that the econometrician has knowledge of past and present public information variables Z_t, but not of Y_t, as well as knowledge of the return series. Also, the econometrician does not observe individual private information signals σ_t^i or labor income y_t^i. On the other hand, the econometrician might use past net flows data to serve as an instrument for variables that might be included in Y_t.

Let $\tilde{\Omega}^t = (Z_t, R_t, \left(Z_{t-h}, R_{t-h}, NF_{t-h}^{US}\right)_{h\geq 1})$ describe the information used by the econometrician to predict net flows. Using $\tilde{\Omega}^t$, measured residuals are

$$\begin{aligned} \mathrm{NF}_t^{US} - E(\mathrm{NF}_t^{US}|\tilde{\Omega}^t) &= \Theta^Y(\Delta Y_t - E[\Delta Y_t|\tilde{\Omega}^t]) \\ &\quad +\Theta^I(\Delta \hat{I}_t - E\left[\Delta \hat{I}_t|\tilde{\Omega}^t\right]) \\ &\quad +\Theta^\sigma\left(\Delta\sigma_t - E\left[\Delta\sigma_t|\tilde{\Omega}^t\right]\right). \end{aligned}$$

Residuals in the net flows regression thus have three components. The first is due to omitted public information variables Y_t. In the econometric analysis below, we address this issue by incorporating a comprehensive list of public information variables in our regressions. The second is that there is payoff relevant information from the past, $\hat{I}_t$, which is not contained in the econometrician's information set. Importantly, to the extent that this information is not in any Z_{t-h} or R_{t-h} for $h > 0$, it might be captured by past net flows, which should thus play an important role in the regression. Indeed, if private information matters, then we would expect $\hat{I}_t$ to move around, and lagged flows to be significant.[15]

Finally, the third component reflects *aggregate private information.* It captures flows due to changes in the aggregate signal that cannot be forecast from public information.[16] We conclude that if the econometrician has the same public information as all investors do, then unexpected net flows

$$\begin{aligned} \tilde{\upsilon}_t &\equiv \mathrm{NF}_t^{US} - E(\mathrm{NF}_t^{US}|\tilde{\Omega}^t) \\ &= \Theta^\sigma\left[\Delta\sigma_t - E\left(\Delta\sigma_t|\tilde{\Omega}^t\right)\right], \end{aligned} \tag{4}$$

provide a measure of aggregate private information of US investors. This private information is not available to any individual investor at the beginning of period t since it relies on knowledge of end-of-period variables. This is true for the econometrician as well. However, an examination of the residuals from a regression of realized net flows on the information set reveals the aggregate private information of US investors.

[15] In our regressions below, the serial correlation of the residuals is very close to zero when lags of the dependent variable are incorporated in the regression. Thus, it seems that assuming that $\hat{I}_{t-1}$ is also observed by the econometrician is a good approximation.

[16] It is easy to show that if signals are uncorrelated with future payoff relevant information, then $\Theta^\sigma = 0$.

The empirical implementation of (4) requires that we take a stand on how periods and trading dates in the model relate to the data we use. To see this, let one period in the model represent one month of calendar time (the frequency of our data). The model assumes that agents trade only on the last day of the month, taking prices as given. Hence, time t variables included in $\tilde{\Omega}^t$ should be understood as end-of-month variables. However, this assumption could lead to a very conservative measure of private information if some private information is revealed by trades that occur during the month.

We therefore construct an alternative measure of private information using beginning-of-period values of the public information variables. Here, let $\Omega^t = \left(Z_{t-h}, R_{t-h}, NF^{US}_{t-h}\right)_{h\geq 1}$ so that no end-of-period variables are included. Using this information set the econometrician will recover the residuals

$$\upsilon_t \equiv \mathrm{NF}^{US}_t - E(\mathrm{NF}^{US}_t|\Omega^t). \tag{5}$$

We label this our "broad" measure of private information. We note that this measure may overstate the effects of private information if investors trade on unexpected public information released during the month. While our "conservative" measure of private information $\tilde{\upsilon}_t$ in (4) accounts for such public information news, it will understate the effects of private information which are incorporated into prices during the period. It turns out that our main results below are quite similar for the two measures.

From our discussion of the optimal asset allocation policy, we see that υ_t and $\tilde{\upsilon}_t$ contain private information relevant to forecasting unexpected returns. We note that both private information measures could be related to hedge demands by the investors which do not impact returns. This is possible if the aggregate of US investors is small enough relative to the market. This is another potential source of noise in υ_t and $\tilde{\upsilon}_t$ and suggests that we might underestimate the impact of private information on returns.

We can use our two measures of private information to examine the structure of the information set used by investors in the international equity markets. It is standard to divide the public information vector Z_t into 'global variables' (such as the US interest rate, which affects equity prices in all of the countries), and 'local variables' (which may affect the equity return in a specific country). For the private information vector υ_t (or $\tilde{\upsilon}_t$) no natural labeling is available. However, suppose there is a linear combination $\phi_t'\upsilon_t$ which explains a large fraction of the variance of υ_t. This suggests that there is a single indicator driving private information which is relevant for trading in *all* markets. In this sense, $\phi'\upsilon_t$ is a 'global factor' in private information. A global factor in private information is thus consistent with traders having private information about the global variables driving factor returns or about a systematic factor in asset prices as in Subrahmanyam (1991).[17]

[17] Work by Chan (1993) shows that common factors in market-wide private information are impor-

We construct such a factor below using factor analysis.

2.3 Private Information and Equity Returns

To test whether our measures of private information impact equity prices we need to specify an estimable model of returns. We adopt the latent-factor model of Hansen and Hodrick (1983) and Gibbons and Ferson (1985). In a $K-$factor model, the market price of risk of the k-th factor can be written as a linear combination of the set of L instruments $Z_t = (Z_{1,t}, ..., Z_{L,t})$ that are in the (public) information set of the econometrician. Thus, the process for returns conditional on the econometrician's information is described by:

$$R_{t+1} = \beta\alpha Z_t + \beta f_{t+1} + \varepsilon_{t+1}. \qquad (6)$$

In this return model, f is a $K \times 1$ vector of factor realizations with $E[f_{t+1}|Z_t] = 0$, $\beta = cov(R_{t+1}, f_{t+1}|Z_t)$ is a constant $N \times K$ matrix, α is a $K \times L$ matrix, and f and ε are uncorrelated. In this model, the linear combination αZ_t represents the latent factors driving returns, while the β matrix is the loading on the factors.[18] We present more details on our empirical implementation of the model below.

There are two reasons for adopting this latent-factor structure. First, the tests for private information effects outlined below require us to evaluate the impact of unexpected flows on unexpected returns. We thus use the "usual suspects" in Z_t which have been shown in many other studies to forecast international equity returns. Second, imposing the over-identifying restrictions of the latent-factor model (detailed below) results in more precise estimates of return variation due to the public and private information variables. This latter effect is important due to the noisiness of monthly return data.[19]

Predicting Returns with Public and Private Information

The latent-factor model of returns summarizes the *public* information relevant for forecasting returns in a parsimonious way. For example, under a one-factor model, the resulting factor αZ_t is often interpreted as a 'global' factor which is relevant for all stock markets (e.g. Campbell and Hamao (1992)). In the same vein, it is interesting to examine whether *private* information is relevant for forecasting returns.

tant in explaining the cross-autocorrelation of domestic stock returns.

[18] The law of motion for returns described in (6) is understood to be consistent with the investor's conditional distribution given by Φ_0, Φ_1, and Σ_{II}.

[19] Note that alternative models for the cross section of expected international equity returns (e.g. the international version of the three-factor model presented in Fama and French (1998)) would use realized returns on sub-portfolios (e.g. international book-to-market portfolios) as proxy variables for the factor returns. These proxies would contain return variation due to the release of private information revealed by the trades of US investors during the month, thus invalidating the tests using our broad measures of private information.

Consider first evaluating the effects of our broad measure of private information on returns. Since Z_{t-1} and υ_t are uncorrelated, we require that

$$E\left[R_t|Z_{t-1},\upsilon_t\right]=\beta\alpha Z_{t-1}+\gamma\upsilon_t.$$

Because of its contemporaneous nature we need to distinguish a positive impact due to private information from one that would result from price pressure due to the unexpected inflow of capital. To overcome this difficulty, we test whether the price impact of the release of private information is long lived in the spirit of Hasbrouck (1991a,b).[20] Let R_t^{t+H} be the cumulative equity return from the beginning of period t to the end of period $t+H$. Hence, we estimate the model:

$$E\left[R_t^{t+H}|Z_{t-1},\upsilon_t\right]=\beta_H\alpha_H Z_{t-1}+\gamma_H\upsilon_t. \tag{7}$$

We discuss sign restrictions of the coefficients below.

Note that our method allows us to keep both the public and private information sets of the investor constant as we increase the holding period. As other investors observe the unanticipated trades of the US investors, they will be able to gain some knowledge of the private information. In effect, some private information will become public information. By holding the public information set constant, we can follow the tests advocated in the microstructure literature and estimate the effects of private information released at time t on future prices (returns) without allowing for this additional contamination.

We also test for the effects of our conservative measure of private information on returns. Consider a regression of next month's returns on the currently observed instruments as well as private information $\tilde{\upsilon}_t$. Since $\tilde{\upsilon}_t$ and Z_t are uncorrelated, we obtain

$$E\left[R_{t+1}|Z_t,\tilde{\upsilon}_t\right]=\beta\alpha Z_t+\widetilde{\gamma}\tilde{\upsilon}_t. \tag{8}$$

The parameter $\widetilde{\gamma}$ measures the contribution of private information to next month's returns.

Based on the portfolio choice model presented above we cannot formally restrict the sign of the coefficients γ_H and $\tilde{\gamma}$. To see this, note that our measure of private information in equation (4) is composed of the signal vector itself times the matrix Θ^{σ}. The elements

[20] In Hasbrouck's microstructure tests, the release of private information has a *permanent* impact on future prices as he assumes expected returns are constant. This is not a very strong assumption as the tests span a short period of calendar time. In his tests, public information is assumed to be random and captured in the residual on the return equation in a vector autoregression. In the monthly observations used in this paper, the assumption of constant expected returns is questionable as returns vary in response to the release of public information. We thus test for the *long-lived* effects of private information on prices after accounting for time-varying expected returns.

of this matrix are a function of the myopic and hedging demands of US investors.[21] We are unable to sign these elements. Thus, in theory, we cannot reject the possibility that a positive private information signal on a country's return leads to an unexpected net outflow from that country.

It is important to distinguish the lack of restrictions on γ_H and $\tilde{\gamma}$ in this model from the approach in standard microstructure models (e.g. Hasbrouck (1991a,b)) where agents are risk neutral. Risk neutrality implies that investors care only about expected returns and have no hedging demands. In these models, unexpected net purchases are associated with positive signals about returns which in turn implies that γ_H and $\tilde{\gamma}$ should be positive. In anticipation of our results below, we find that the coefficients are mostly positive suggesting that hedging demands have a small impact on trades by US investors.

We use the same method to test the impact of the global factor on the cross section of returns. However, note that there is no reason to believe that the global factor impacts returns in each individual country. This is because we construct the global factor using factor analysis which only requires that the linear combination $\phi_t' \upsilon_t$ is composed of some, but not all, of the elements in υ_t. As we stated in the introduction, if the global factor in private information is related to the global stock market return, then a sufficient test is the joint significance of the factor across all of the returns.

Finally, the model places no restrictions on the absolute magnitude of the impact on individual country returns from each measure of private information. Below, we pursue the estimation of the measures of private information (υ_t and $\tilde{\upsilon}_t$), their effects on the cross section of equity returns (γ_H and $\widetilde{\gamma}$) and the global factors in each of the measures of private information. Inference on these parameters is at the center of the empirical tests in the paper.

3 Modeling Expected International Flow and Return Variation

We separate our empirical analysis of the cross section of expected international equity flows and returns into three parts. In the first part, we briefly describe the data. In the second part, we provide further details on the empirical model of international equity flows introduced in section 2.2 and describe how we construct our measures of private information. In the third part, we present the latent-variable model of the cross section of international equity returns.

[21] In the appendix we show that in this model, as in other portfolio choice models, asset demands have two components: a myopic demand and an hedging demand. We also show that both of these components are linear in I_t so that the matrix Θ^σ results from a combination of the two.

3.1 Data

Our data on the international equity flows of US investors is obtained from the Treasury International Capital (TIC) reporting system of the US Treasury.[22] Financial institutions (banks, bank holding companies, securities brokers, dealers, and non-banking enterprises) must report to the Treasury each month on all of their transactions with foreigners in long-term securities (e.g. stocks and bonds) by country if their aggregate purchases or sales total more than US $2 million in the month. As a result, the Treasury receives comprehensive data on cross-border equity transactions for most US investors.

In this paper, we examine transactions by US investors in the equity markets of eight large developed countries – Germany, Japan, UK, France, Canada, Netherlands, Switzerland and Italy – which account for approximately 74 per cent of the market value of non-US markets at the end of our sample period. There are several advantages to limiting our analysis to these countries. First, using a group of relatively homogeneous, developed countries allows us to measure the public and private components of the trades correctly. Flows in and out of the equity markets in these countries are likely to be driven by stable economic relationships for which there is an abundant list of instruments. In contrast, the on-going process of liberalization of equity markets in developing countries leads to capital flows that are mostly driven by changing risk-sharing opportunities or declining transactions costs (e.g. Stulz (1999)). Distinguishing this effect from the asymmetric information effect in the flows would be difficult. Similarly, the assumption of no trading frictions used in the model is more at odds with flows into emerging markets.

Second, as a result of the liberalizations, models of portfolio flows into emerging markets suffer from several statistical problems that make them difficult to estimate (e.g. non-stationarity due to structural breaks from changing foreign ownership restrictions). This is perhaps one of the reasons why researchers have focused on aggregated/regional data across these countries (e.g. Froot et al. (2001), and Bekaert et al. (1999)). Third, our data contain the net purchases of equities across a large cross section of countries with long, coincident time spans. This is typically not available for microstructure data sets and is clearly necessary in order to measure common components in unexpected flows and their effects on asset prices.

However, there are two main deficiencies with the data. First, the Treasury does not collect data on transactions in derivative securities which have grown in importance in recent years. For example, investors may decide to take a position in an international

[22] There are a number of related studies that use the same data set (Tesar and Werner (1993, 1995), Bohn and Tesar (1996a,b) and Brennan and Cao (1997)). See Froot et al. (2001) and Levich (1994) for a description of limitations/advantages of US Treasury data.

equity market by purchasing a stock index futures contract at a lower transaction cost than investing directly in the equity market. To the best of our knowledge, this criticism applies to all datasets used in this literature. Second, and perhaps more importantly, the Treasury collects data by geographic center and not by the country of origin of the security. This means that the data may be unrepresentative for those countries that contain large international financial centers such as the UK and Switzerland. The typical example of this is a European company that is issuing securities in the Euro-equity market and sells the securities through banks in London to US investors. This transaction would be recorded as a sale of UK equity. Warnock and Cleaver (2002) examine the TIC data in detail and find that transactions to the UK are overstated while transaction to other countries are understated. This may bias our results against finding significant private information effects as the companies that are likely to issue equities via the Euro-equity market are large companies that are well known outside their home country (Marr et al. (1991)). However, these are the companies that foreign investors tend to hold, as noted above. Thus, the TIC would seem to capture disproportionately net flows for those companies that are less likely to be affected by global components of information. [23]

The flows data all exhibit a trend growth as US investors (slowly) increase their net holdings of foreign equities. We thus follow the standard approach in the literature and normalize the flows data by dividing by the beginning-of-period value of the foreign equity market index.[24] The normalized net flows are also a better measure to describe the behavior of the investors in our model (recall that NF_t^{US} is measured in number of shares). Table 1a details our data set which ranges from January, 1977 (the start of the monthly TIC data) to September, 2000. The data are separated into gross purchases, gross sales and net purchases of foreign equities by US residents. The volume of transactions in the gross flows are much larger than in the net flows. US investors have purchased (sold) foreign equities in amounts ranging from 0.135 per cent (0.121 per cent) per month of the market value of Japanese equities up to 0.718 per cent (0.681 per cent) per month of UK equities. The low volume of transaction in the Japanese equities is likely related to the poor performance of the Japanese equity market throughout the 1990s. The high volume of transactions in UK securities is likely overstated due to the presence of London as an international financial center. In contrast, there is a much smaller volume of net purchases which range from 0.011 per cent per month for German equities up to 0.055 per cent per month for Canadian equities. The gross flows have higher volatilities than

[23] In addition, random measurement errors in the data will reduce the significance of any measures of private information. In our analysis below, when we exclude the UK data our results are strengthened.

[24] We use the Datastream total market index for the country's equity market and convert it to US dollars.

the net flows as well as higher autocorrelations. However, all of the normalized series appear stationary.

We use the Datastream equity return index available at the end of each month to construct monthly returns. The indexes are based on the firms with the largest equity values in each country. As there is evidence that foreign investors tilt their portfolios towards the largest companies in each market, these indexes should be close to those returns actually obtained by US investors in foreign markets. The returns are translated into US dollars and the risk-free US interest rate is subtracted from them. Table 1b presents the summary statistics of the excess returns. We note that the volatility of the returns is much larger than that of the flows. This has implications below when we try to model the expected portions of flows and returns.

Portfolio flows and equity returns are contemporaneously correlated. Table 1c shows the correlation coefficients between the gross purchases, gross sales, and net purchases by US investors and the contemporaneous excess equity return in the country. Gross purchases have correlation coefficients of 0.10 or more with the foreign stock return for all countries except the UK (0.003) and Switzerland (0.022). The correlations on gross sales are lower, with the figures for the UK and Switzerland being negative. The correlation between net purchases and equity returns are larger than 0.10 for all countries except the Netherlands (0.092) and the UK and Switzerland (both less than 0.050). This suggests either that US investors allocate funds to these countries when expected returns are high and/or that the purchases themselves drive up prices. Our test method presented in section 2.3 allows us to distinguish between these effects. The coefficients also suggest that the large volume of purchases and sales recorded for the UK and Switzerland are more likely due to activity in the Euro-equity market.

3.2 Modeling Expected International Equity Flows

3.2.1 Public information sets

To estimate private information in the international equity markets, we need to separate the equity flows into expected and unexpected components. We use four different sets of public information variables to show that our results do not depend on which variables we choose.

In the first public information set (A), we use an autoregressive model to capture the expected portion of the flows. As noted in Table 1a, the flows are persistent and a third-order autoregressive model is able to capture a substantial portion of the expected flows. This is similar to the approach by Warther (1995) who uses autoregressive models for mutual fund flows.[25]

In the second and third information sets (B and C), we choose public information variables based on the belief that monthly international capital flows are related to the cross section of expected international returns as in our model in section 2.[26] The persistence in the flows is also consistent with the existing evidence that expected equity returns are time varying and persistent (see Chapter 7 in Campbell et al. (1997)). Hence, to model the expected portion of flows, we will use instruments that have been shown to predict the cross section of international equity returns.

In this literature, it is common to separate the instruments into two groups, global and local (e.g. Harvey (1991), Ferson and Harvey (1993,1994)), where the use of local variables is justified by the existence of incomplete risk sharing. From this literature we select the short-term US interest rate, the US credit spread, the dividend yield on the global stock market, and the slope of the US term structure to act as "global" variables. These variables can be thought of as capturing shocks to wealth and risk tolerance of US investors and are lagged by one period so they are observable by investors before any portfolio decisions would be made. The global variables, along with lagged flow variables to capture any missing expected flows, are used as our second data set (B). Table 1b provides the summary statistics of the global instruments, which resemble those in many other papers.[27]

We also select a number of "local" variables for each foreign country. These include: the lagged value of the country's stock return (to capture any return chasing activities by investors (Bohn and Tesar (1996a,b), Froot, O'Connell, Seasholes (2001)); the lagged value of the country's dividend yield and the difference in interest rates between the country and the US (to act as instruments for expected return variation specific to the country); and last month's change in the spot exchange rate of the country's currency against the US dollar (to capture any exchange rate effects). The third information set

[25] We have tried extending the lag lengths of our empirical flow models out to a six-month lag. Our tests are robust to this choice (results available on request).

[26] Bohn and Tesar (1996a,b) find that net purchases are related to the cross section of expected equity returns using a similar list of instruments. They also separate their instruments into local and global groups. Our approach differs by using the variables directly in modeling gross purchases and sales, which turn out to be much more predictable than net purchases.

[27] Most of the data are from Datastream. We use the 30-day Eurocurrency rate on the last business day of each month as the risk-free interest rate for each country. We use the Datastream equity return index and dividend yield on the index available at the end of each month. The dividend yield series is deseasonalized in the usual manner (we take the average dividend over the last 12 months and divide by the current price). The exchange rate is reported daily by Datastream in UK pounds and we translate the end-of-month values into US dollar equivalents. The US credit spread is the difference between the AAA and BBB bond yields available at the US Federal Reserve web site. The term structure slope is the long bond yield from the OECD database less the 30 day Eurodollar interest rate.

(C) thus contains lagged flows, and the global and local information variables.

To construct our fourth information set (D), we rely on the literature that examines the links between trading activity and market volatility. This literature shows a positive contemporaneous correlation between conditional volatility and trading volume (e.g. Lamoureux and Lastrapes (1990), Gallant et al. (1992)). As our international equity flows represent the trading activity of a specific group of investors, it may be that they are correlated with conditional volatility and that instruments which forecast time-varying second moments can also forecast the flows. We thus use a very simple autoregressive conditional heteroskedasticity (ARCH) type representation to capture these potential effects. We form a small vector autoregression (VAR of order 1) composed of the US excess equity return, the foreign country excess equity return and the change in the spot exchange rate between the two countries. The squared and lagged estimated residuals from this VAR act as instruments to forecast future flows. Again we include lagged flows to capture any of the missing variation.

3.2.2 Regression results

Summary statistics for OLS regressions of gross purchases and gross sales on the four different information sets are presented in Tables 2a and 2b, respectively. These regressions use beginning-of-month values of the variables. The regressions using the end-of-month values produce similar results and are available on request. A simple AR(3) specification (information set A) is able to capture a large part of the expected flow variation with $\bar{R}^2$ measures (adjusted for degrees of freedom) ranging from 0.314 for France up to 0.933 for the UK for gross purchases by US residents. A similar range is recorded for gross sales by US residents. The combination of the lagged flow variables and the global instruments (information set B) also results in very high $\bar{R}^2$ measures for the gross flows. When we test the joint significance of the global variables (excluding lagged flows) we find that the restriction that all of the coefficients on the global instruments are zero is strongly rejected by the data for most countries. Thus, the global instruments that are typically used to explain the cross section of international equity returns also explain the cross section of international equity flows. It is clear, however, that most of the predictability comes from including the lagged values of the flows as the $\bar{R}^2$ statistics increase only marginally by including the global variables.

We test for the predictability of the local instruments by adding them to the existing regression (information set C). Table 2 shows the resulting $\bar{R}^2$ measures and the χ^2 tests of their joint significance. There is a small increase in the $\bar{R}^2$ statistics from adding the local variables to the regression except for those explaining gross purchases of German and Italian equities and gross sales of French, Canadian and Italian equities. However, the local variables as a group are jointly significant at the 5 per cent level for equity flows in and out of all countries except for gross purchases of German and French stocks and gross sales of German and Canadian stocks.

The fourth information set (D) uses lagged values of squared residuals to capture any time variation in flows related to time-varying volatility. The statistics show that these instruments do not help to explain time-variation in expected flows beyond that captured by the lagged flow itself.

Most of the existing literature has focussed on explaining the expected component of *net* portfolio equity flows. While we find that the gross flows are quite predictable, the net flows are much less so. Table 2c provides the $\bar{R}^2$ and χ^2 statistics for the OLS regressions of the net equity flows on the four information sets. As can be seen, the amount of linear predictability in the net flows is much smaller than in the gross flows. The adjusted $\bar{R}^2$ statistics for the net flow regressions are below their counterparts for the gross flow regressions, with the former ranging from 0.021 for Italy to 0.338 for Japan using the three lagged net flows as regressors (instrument set A). The global, local and heteroskedastic variables also do not appear to capture much of the predictability in the net flows regressions.[28]

It appears that much of the predictability in our gross flow regressions comes from modeling the autocorrelation structure of the data. Bohn and Tesar (1996a) note the strong autocorrelation in the *net* purchases data. They view this as a challenge for portfolio theory as investors should adjust their portfolios in response to news about expected returns which have little serial correlation. In the model this is justified by assuming that lagged flows are good proxies for past information not in the econometrician's information set. In general, there are other economic reasons to believe that flows exhibit "momentum" (Bohn and Tesar (1996a)). First, financial firms may set long-run asset holding targets that lead to smoothing and slowly trending flows. Second, herding by US investors at low frequencies would also deliver autocorrelation of flows.

Our results also have interest for more general international asset pricing applications that use similar forecasting variables. One potential criticism of most asset pricing applications is that many of these variables have been chosen by an on-going implicit process of data snooping; i.e. choosing the variables based on ex-post statistical criteria of return

[28] The small amount of predictability due to the global and local instruments for both the gross and net flows is somewhat surprising given the relationships found in the domestic market literature. For example, Hasbrouck and Seppi (2001) find that statistical factors in the cross section of order flows are related to those in the cross section of returns on the 30 Dow Jones stocks. Lo and Wang (2000) find that some of the standard instruments used to predict stock market returns help predict turnover of the NYSE-AMEX stocks.

predictability. To the extent that the flows data represent new sources of information (given the correlations in Table 1c), the (limited) predictability shown here may alleviate some of these concerns.

3.2.3 Constructing measures of private information of US investors

The regressions detailed in the previous section give us a way of estimating the expected portion of the gross and net portfolio equity flows. Following our derivations in section 2, the private information of the US residents can be estimated by the unexpected portion of the net equity flows.[29] In this subsection we discuss only the construction of our broad measure of private information, and note that the exact same steps apply for our conservative measure.

Recall from equation (5) that the vector of our broad measures of private information is obtained from the unexpected net portfolio flows into the equity markets of the eight foreign countries and is denoted by $\upsilon_t = (\upsilon_{1,t}, \ldots, \upsilon_{8,t})$. Our regression results allow two possible routes to estimating υ_t. One is to use net purchases of equities by US investors in each foreign country as the dependent variable (Table 2c). The residual from that regression is υ_t as desired. An alternative route is to estimate the regressions for gross sales and gross purchases separately (Tables 2a and 2b) and let

$$\upsilon_{n,t} = \upsilon^P_{n,t} - \upsilon^S_{n,t}, \tag{9}$$

where $\upsilon^P_{n,t}$ and $\upsilon^S_{n,t}$ are the residuals from the gross purchases and gross sales regressions, respectively.

Although the gross flow regressions show substantially greater explanatory power than their net flow counterparts, it is uncertain as to whether the expected portion of the net flows are modelled better by using the difference between the expected gross flows or by using the net flows directly in a regression. To answer this, we can construct an expected net flow variance ratio. The numerator is the variance of the expected net flow from the net flow regressions (Table 2c). The denominator is the variance of the implied expected net flow constructed using the difference between the expected gross purchases and expected gross sales (Tables 2a,b). A ratio below 1.0 indicates that the approach using the gross flow regressions is to be preferred. We can also test whether this difference is statistically significant. We regress the net flows on both the lagged gross purchases and gross sales and test if the coefficients on the two sets of variables

[29] Kaufmann, Mehrez, and Schmukler (1999) proceed in a similar fashion to obtain a measure of private information. They use survey data on firm managers' assessments of how the economy will preform. The advantage of using flow data is that investors actually "put their money where their mouth is."

are equal (and opposite in sign). The resulting test statistic is χ^2 distributed with three degrees of freedom (using the third-order autoregressive model).

Table 3 presents the estimated variance ratios and the asymptotic marginal significance levels (P-values) of the test statistics. For all countries and information sets, the ratios are below 1.0, often by a substantial amount. It appears that this result is primarily driven by the different time-series properties of the gross purchases and sales which are obscured when one models net flows directly. This is shown in the first column of the table which compares the variances of the estimates using the AR models for the flows. Adding on other information variables to the basic AR specification raises the variance ratios in all cases. The P-values indicate that the differences between the coefficients on the lagged purchases and sales are statistically different at the 10 per cent level in five of the eight countries examined for instrument set C.

To conclude, using the residuals from the gross flow regressions (9) to obtain our broad measure of private information is our preferred approach. This way, as Table 3 shows, we are able to extract more public information out of net flows. Similar results are obtained when constructing our conservative measure of private information.

3.3 The Baseline Model of International Equity Returns

3.3.1 The latent-variable model of international stock returns

The procedure outlined above allows us to separate the equity flows into expected and unexpected components. The unexpected component of the net flow is a measure of the private information in these markets and our task is to see if it affects realized equity returns (see subsection 2.2). To do this we use the baseline empirical asset pricing model presented in section 2 which we repeat for convenience:

$$E\left[R_t^{t+H}|Z_{t-1}, \upsilon_t\right] = \beta_H \alpha_H Z_{t-1} + \gamma_H \upsilon_t, \tag{10}$$

for the broad measure of private information and

$$E\left[R_{t+1}|Z_t, \tilde{\upsilon}_t\right] = \beta \alpha Z_t + \widetilde{\gamma} \tilde{\upsilon}_t, \tag{11}$$

for the conservative measure of private information.

Recall that Z_t represents the public information variables observed by the econometrician at the start of the month. The typical element of the coefficient matrix $\Phi_H = \beta_H \alpha_H$ is $\Phi_{H,n,l} = \sum_{k=1}^{K} \beta_{H,n,k} \alpha_{H,l,k}$, $n = 1, ..., N$ and $l = 1, ..., L$. The α_H coefficients capture the effects of time-varying returns on the latent factors over the holding period H and the β_H coefficients capture how the cross section of international returns respond to the factor returns. We extend the holding period H of the investor from $H = 0$, where the stocks are held for the current month, out to three months ($H = 3$). The model is estimated separately for each holding period H.[30]

The model is estimated by Generalized Method of Moments (GMM). We follow the standard practice for identification and assume $\beta_1 = 1$. The model imposes cross-equation restrictions so that Φ_H has a reduced rank structure with $\text{rank}(\Phi_H) = K$. These restrictions can be used as a test of the model using the GMM J-statistic which is distributed as χ^2 with $(N - K)(L - K)$ degrees of freedom.[31]

Latent-variable models of international stock returns have been used by Harvey (1991), Campbell and Hamao (1992), and Harvey, Solnik and Zhou (1994) among others. The results of these studies are mixed. Campbell and Hamao (1992) examine the integration of the US and Japanese equity markets and find that a single latent-variable model is rejected during the 1970s but not during the 1980s. Harvey (1991) finds that the data reject a single source of risk across all of the world's equity markets, implying that the world market portfolio is not conditionally mean-variance efficient. However, the rejection is strongest for Japan, while he finds the model holds for the other countries examined in the paper. Harvey et al. (1994) find that a one to three latent-factor model is rejected by the cross section of 18 country index returns. However, when they examine the models' pricing errors and variance ratios, they find that a two or three latent-variable model captures the cross section of country returns.[32]

[30] Cumby (1989) and Lewis (1990,1991) perform similar tests of longer holding periods in latent factor models.

[31] We use the Newey-West form of the asymptotic covariance matrix with the number of lags equal to $H + 3$ to capture any residual autocorrelation. Ferson and Foerster (1994) examine the small-sample properties of latent-variable models estimated by GMM. They find that an iterated GMM procedure results in coefficient estimates with small biases but that the standard errors are understated. They propose a correction factor which results in appropriate sized standard errors. We use the iterated GMM approach and apply their small-sample correction factor to our standard errors.

[32] One of the key assumptions in the latent variable model is that the betas are constant. This not a strong assumption at the country level. Ferson and Harvey (1993) test an asset pricing model where the risk factors are global but the conditional betas depend on country-specific attributes. They are interested in the portion of variation of country returns that can be captured by time-varying risk premia as opposed to time-varying betas. They find that although time variation in the betas is statistically significant it contributes little to the variation in expected returns. Ferson and Harvey (1994) examine whether country-specific fundamental attributes can be used to help motivate time-varying beta models. Again, the risk premia are global while the betas are functions of specific country attributes which they label "fundamental determinants". They find some limited support for their model. However, the estimation approach used in both of these papers does not allow the cross-equation restrictions of a global asset pricing model to be imposed.

3.3.2 Results

To show that our global (public information) instrument set is able to capture time-varying expected returns, we follow the literature and present some initial OLS regressions of the excess stock returns on the global instruments without imposing any of the restrictions of the asset pricing model or including the effects of private information. Table 4 presents the $\bar{R}^2$ measures (adjusted for degrees of freedom) as well as the χ^2 statistic of the test that the global variables are jointly insignificant. The low value of the $\bar{R}^2$ statistics is typical in the equity return prediction literature. The χ^2 statistics do show however, that the global variables are significant at the 5 per cent level for 4 of the 9 countries (Germany, Japan, Netherlands and Switzerland) and significant at the 10 per cent level for the US. The instruments are not significant at standard levels for the UK, French, Canadian or Italian equity returns.

We also test whether the same local instruments that were used in the flow regressions also are able to capture expected return variation. When these variables are included there is only a small increase in the degree of linear predictability in some of the countries (Table 4). In addition, the χ^2 statistics show that the local variables are significant at the 10 per cent level for the UK only (P-value of 0.088). Thus, it appears that the cross-section of international equity returns can be modelled by the global instrument set alone.[33]

Imposing the restrictions of the latent-variable model leads to more precise estimates of the coefficients on the global information variables. Table 5a presents the estimated α coefficients for holding periods ranging from the current month ($H = 0$) to three months forward ($H = 3$). The coefficients are statistically significant on all of the variables (except the intercept) at all forecast horizons. The coefficient on the short-term US interest rate is negative as has been shown in other studies. The coefficients on the other three variables are positive. The estimated β coefficients are presented in Table 5b. The coefficients on the foreign countries are also precisely estimated with all of the coefficients significant at conventional levels. The coefficients are all around 1 at the one month holding period ($H = 0$) and thus seem reasonable from an economic point of view. For longer holding periods, the Japanese coefficients are high (near 2) while the German and UK coefficients are lower. This may be indicating either that these equity returns

[33] This is a different picture than the one given for flows where local factors were shown to (marginally) increase the models' explanatory power. If these local variables represent the effects of a local factor in flows then it may be that there is also a local factor in returns. However, the return regressions may not capture this factor as returns are much noisier than flows. This does not affect our tests below as the measures of private information constructed using instrument set C are orthogonal to these variables.

respond differently to the global risk factor as the forecast horizon lengthens or that the model may be missing an additional risk factor.[34]

The model is also able to capture return variation due to public information. To show this, we construct a variance-ratio statistic similar to the ones presented by Campbell and Hamao (1992) and Harvey (1991). The numerator is the variance of the fitted values from equation (10) denoted $var(\hat{\beta}_H \hat{\alpha}_H Z_{t-1})$ where $\hat{\beta}_H$ and $\hat{\alpha}_H$ are the GMM coefficients for holding period H given in Tables 5a and 5b, respectively. The denominator is the variance of the fitted values from an OLS regression of the excess return on the global instruments $var(\widehat{\delta} Z_{t-1})$. The variance ratio thus shows how imposing the baseline model's over-identifying restrictions leads to a degradation of the data's ability to forecast expected returns.

The results are presented in Table 5c. As can be seen, the latent-variable model does a good job at capturing expected return variation. The ratios are above 1.00 for the current month ($H = 0$) for the returns on equities from six of the nine countries (German (0.971), UK (0.940) and French (0.712) stock returns are the exceptions). As the holding period lengthens, the impact of the model's restrictions increase and some of the ratios fall below 1.00. However, even with a three-month holding period ($H = 3$), five of the nine ratios are above 1.00.

In Table 5c, we also present the J-statistics that evaluate the over-identifying restrictions of the latent-variable model. The J-statistics show that the model is not rejected at any forecast horizon. Given the good performance of the one-factor latent-variable model according to all of these metrics, we will use it as our base model for subsequent tests of private information in international markets.

4 Private Information Test Results

Equipped with our measures of US investors' private information and a baseline model of expected return variation, we now go on to analyze whether the release of private information affects returns. We start by presenting our results for the measures of total private information. We then use a factor analysis on the covariance matrix of private information to identify global components in the two measures of private information. Finally, we show how our results are robust to alternative assumptions.

[34] We note below that our results are robust to missing risk factors in returns.

4.1 Total Private Information Test Results

The top panel of Table 6 presents the value of the chi-squared statistics associated with the Wald test of the null hypothesis that all of the γ_H coefficients in (10) are jointly equal to zero.[35] The broad measures of private information are jointly significant across all of the foreign countries and holding periods regardless of the instrument set used to construct the expected equity flows with the P-value on each of the test statistics less than 0.001. We thus find a significant role for private information effects in the cross section of international stock returns.

The left panel in Table 7 presents the individual γ_H coefficients on the broad measures of private information over the different holding periods using the global and local information variables in the flow regressions (instrument set C). Each coefficient is presented with the small-sample standard error and P-value. The joint significance tests presented in Table 6 are repeated here for convenience. The coefficients are positive for all of the returns in the current month ($H = 0$) indicating that the private information available to US investors leads to higher stock returns. As US residents make an unexpected net purchase of the stocks in the foreign country, the foreign stock index rises. As discussed before, this suggests that hedging demands of US investors are small.

The magnitude of the coefficients is in basis points. Thus a one basis point increase in the (normalized value of) US net purchases of German equities leads to a 12.8 basis point increase in German excess equity returns (in US dollars). Below, we present some variance-ratio statistics to get a better measure of the economic importance of private information on returns. The individual coefficients are significant at the 10 per cent level for all but two of the equity returns (UK and Italian stocks are the exceptions).

The release of private information has a long-run *price* impact as the coefficients over longer holding periods ($H = 1$ to $H = 3$) are also mostly positive and statistically significant. Some of the estimated coefficients measuring the effects of private information on UK and Swiss equity returns are poorly estimated due to these countries acting as financial centers. Overall, however, the combined results of Tables 6 (top panel) and 7 reveal that the unexpected component of US residents net purchases of foreign securities leads to a long-run increase in the prices on the stocks. We attribute this to the private information effect and discuss several potential alternative explanations below.

The right panel of Table 7 presents the results for our conservative measure of private

[35] Some authors (e.g. Harvey and Zhou (1993)) have noted potential problems with Wald tests in tests in systems with instrumental variables and have advocated calculating the Gallant-Jorgensen (1979) (G-J) test statistic as well. We have calculated the G-J test statistics and found that all of the G-J statistics are larger in value than our reported Wald statistics. Both test statistics are χ^2 distributed with 8 degrees of freedom.

information constructed by including values of the variables available at the end of the month in our empirical model of the flows. The chi-squared statistic associated with the Wald test of the null hypothesis that all of the $\widetilde{\gamma}$ coefficients are jointly equal to zero is presented at the bottom of the table. With a P-value of less than 0.001 there is a clear effect of private information released this month on the returns in the subsequent month. The coefficients on the individual country returns are naturally not as large nor significant as their counterparts in the left panel of Table 7. However, only one of the coefficients is statistically significantly negative (Netherlands).

Other papers that examine high-frequency data on international portfolio transactions also find that there is a persistent impact of unexpected net flows on returns. Froot et al. (2001) estimate a VAR on flows and returns similar to that of Hasbrouck (1991a) and find persistent price impacts from unexpected flows that last for at least 60 days.[36] However, they are unable to determine whether this low-frequency predictability is caused by the private information of US investors or by price pressure. In a subsequent analysis, Froot and Ramadorai (2001) use closed-end country fund data to separate the effects of private information from price pressure. They find evidence in favor of the information story with US investor flows resulting in persistent price impacts that last for several weeks. However, as they note, using such an approach depends on the adequacy of the model of closed-end fund demand.

By relying on a longer data set and an asset-pricing model, our approach examines the low-frequency predictability more directly. In particular, we can account for the autocorrelation in the gross flows as well as the effects of public information releases on both flows and returns. The results show a significant impact of total private information on international equity returns. We note that this information is superior relative to *all* non-US investors; US investors may not have superior information about foreign equity markets relative to the investors in that particular country.

4.2 Global Factors in Private Information

4.2.1 Extracting common factors in private Information

We now turn to a further examination of the properties of our measure of private information υ_t. We use the variance-covariance matrix of υ_t to inquire about the sources of private information. Given the large degree of predictability in the gross flow regressions, we perform a factor analysis on the residuals from the gross purchases and sales regressions separately ($\upsilon^P_{n,t}$ and $\upsilon^S_{n,t}$, $n = 1, ..., N$). The factor analysis is done by both the

[36] In their structural VAR, Froot et al. (2001) find no impact from unexpected flows on subsequent returns in developed countries.

method of iterated principal factors and by maximum-likelihood estimation. The results presented below use the maximum-likelihood results, but are very similar across the two methods.

Table 8 presents the results of the factor analysis on the residuals from the gross flow regressions using the lagged flows, global and local variables (instrument set C) for the broad measure of private information. The results for the conservative measure are similar and available on request. The first factor captures 56.9 per cent of the unanticipated purchases and 53.3 per cent of the unanticipated sales. Adding in the next two factors raises the total variation captured to over 90 per cent of gross purchases and over 80 per cent of gross sales. We can perform likelihood-ratio tests for the number of factors using the results of the maximum-likelihood factor method. The tests reject the hypothesis that no factors are present in the residuals but give conflicting results about the precise number of factors. The gross purchases residuals reject a one-factor representation in favor of two factors with a P-value of 0.001 per cent. However, a representation involving more than one factor model is rejected for the gross sales residuals. Tests reject more than two factors for both sets of residuals (not reported). Clearly then, the covariance matrices of the gross flow residuals show reduced rank structures associated with a common factor representation in total private information.[37]

To measure the impact of global information in net flows, we obtain the factors estimated using the maximum likelihood method on the gross purchases and sales residuals. As the scale of the factors is arbitrary, we normalize each to have a standard deviation equal to the simple average standard deviation of its constituent residuals. We can then obtain a measure of the global factor in net flows as

$$\upsilon_{n,t}^{G} = \upsilon_{n,t}^{P,G} - \upsilon_{n,t}^{S,G}, \tag{12}$$

where $\upsilon_{n,t}^{P,G}$ and $\upsilon_{n,t}^{S,G}$ are the (normalized) global factors in the residuals from the gross purchases and sales residuals, respectively.

4.2.2 Effects of common factors in private information

If the global factor in private information is related to the global stock market return, then a sufficient test is the joint significance of the factor across all of the returns. We use the same systems of test equations (10) and (11) to perform these tests. A further requirement that the global factor is related to asymmetric information on global stock

[37] Froot et al. (2001) showed in the earlier versions of the paper that there is an important regional factor in *expected* flows. We have removed the expected portion of the flows using the global and local instrument sets and our factor analysis is thus concerned with *unexpected* flows.

returns is that it helps forecast US equity returns. We thus include the global factor in the US equity return regression.

The bottom panel of Table 6 presents the values the chi-squared test statistics associated with the Wald tests that the γ_H coefficients on the global factor in the broad measure of private information are jointly equal to zero. The test statistics are χ^2distributed with nine degrees of freedom. As can be seen, the global factor in private information is jointly significant across all of the countries' equity returns with a P-value of 0.02 or less for the first three information sets. We thus find a significant impact of the global component to the broad measure of private information in our sample.

The left panel in Table 9 presents the individual γ_H coefficients measuring the impact of the global factor in our broad measure of private information on the individual country equity returns along with their small-sample standard errors and marginal significance levels. The coefficients are mostly positive, indicating that as US investors make unexpected purchases across all of the foreign countries the foreign equity indexes rise in response to the release of private information. As before, this is suggestive that US investors on aggregate have small hedging demands. We note that the global factor is positive and significant for US equity returns even though the factor was constructed from net purchases in foreign equity markets. The fact that the global factor is positively related to both US and foreign equity returns indicates that it is related to the common component present in world-wide equity returns.

As the global factor is estimated from a small cross section (eight) of the individual country unexpected flows it is a relatively noisier variable and the individual significance of the coefficients are lower than their total effect counterparts in Table 7.[38] The coefficients reveal that the global factor is significant at the 5 per cent level for all of the countries except Germany, Japan and Italy during the current month ($H = 0$). As the holding period lengthens, the coefficients remain mostly positive, though their individual statistical significance is reduced in some cases.

We have also constructed an estimate of the global factor using our more conservative measure of private information. The right panel of Table 9 shows the results. The chi-squared statistic at the bottom of the table shows that there is statistically significant predictability of the common component in unanticipated net flows this month for the cross section of equity returns in the subsequent month.

Recall that our conservative measure of private information uses the end-of-month values of all of the forecasting variables to remove any potential simultaneity problems

[38] Excluding the UK and Swiss unexpected flows (which are quite noisy due to the Euro equity market activity mentioned above) from the construction of the global factor improves the individual statistical significance measures.

where flows respond to realized returns or other public information released during the month. Thus, our results that show an effect of the global measure of private information of US investors on international equity returns are also free of any potential bias due to contemporaneous trend following of US investors in many foreign markets. However, for the same reason, the conservative measure will not capture a lot of the private information released during the month. As a result, obtaining precise estimates of the individual country's loadings on the global factor in private information will be more difficult. The right panel of Table 9 bears out this intuition where it is shown that the individual country coefficients are imprecisely estimated although they are jointly significant.

4.3 Interpretation

Our results showing a global factor in the private information sets of US investors are in contrast with the perceived notion in the academic literature that investors in international equity markets trade mostly on specific knowledge about certain countries. The most likely interpretation of a common factor in US investor private information is that it originates in superior information about US economic variables. This superior information may come from several sources. For example, US investors could have superior knowledge about US business cycle variation. Lumsdaine and Prasad (1999) estimate a common component in industrial production growth rates in seventeen OECD countries. The average weight of US industrial production in this common component is 38.92 per cent, well above the weights for all of the other countries. This common component is shown to drive industrial production in all of the countries. Kwark (1999) finds that shocks to US output are important in explaining shocks to foreign country output. Superior knowledge of US industrial production or aggregate output would then be of benefit in many countries.

US investors could also have a superior ability to interpret US monetary or productivity shocks which would help in forecasting growth in foreign economies. Kim (2001) estimates several different identified vector autoregression models of the effects of US monetary policy shocks on the current account and foreign country growth. He finds that expansionary shocks to US monetary policy increase imports in the short run. It also causes higher growth in the non-US developed countries with much of the effect coming from a reduced real world interest rate. Glick and Rogoff (1995) analyze productivity shocks in the manufacturing industries of seven major industrial countries and find that gross investment responds to a global productivity shock.

If the equity flows in our data set are predominately those of sophisticated large US investors (e.g. international equity mutual funds and hedge funds) that are active traders in foreign markets, then monitoring the customer order flows of these investors into these particular markets would help more sophisticated investors. US investors may also have private knowledge about links between US firms and firms in particular foreign countries. This may arise from customer or supplier relationships or proposed merger or acquisition activity. Gehrig (1998) discusses the information externalities that are present in large financial centers such as New York.

As discussed in the introduction, the finding of a global factor in private information may help explain the foreign asset holdings of US investors. Kang and Stulz (1997), Choe, Kho, and Stulz (2001), and Dahlquist and Robertson (2001) show that foreign investors tilt their foreign portfolios towards large firms or firms with significant foreign exposures. Presumably these firms have greater exposure to global factors in public and private information. Our finding is also consistent with those papers that show that foreign investors do not underperform domestic investors (Froot and Ramadorai (2001), Choe, Kho, and Stulz (2001)). It is also consistent with the results in Tkac (2001) who examines the performance of US managers of international open-end mutual funds during the 1990-1999 period. She finds that managers of well-diversified funds have significant positive Jensen alphas relative to a global equity index. In contrast, most managers of regional or country-specific funds do not outperform the appropriate benchmark fund. These results are to be expected as the global factor of private information will be relatively more important for well-diversified international mutual funds.

However, one issue that our analysis cannot address is the relevance of the global factor in private information across investors from different countries. For example, it may be that French investors have superior total private information about French securities as they have better country-specific information while having worse global private information. Hence, for these investors, the global factor in private information constitutes a smaller component of total private information.

4.4 Potential Alternative Explanations

Our measures of private information may be contentious as we have extracted them from monthly data and there are a number of potential alternative explanations of our findings. One potential alternative explanation is price pressure: unanticipated net inflows into a country's equity market cause prices to rise even if they are devoid of information content as the market absorbs the extra demand. However, there are two primary reasons why the price pressure explanation is unlikely using our monthly data. First, the net flows shown in Table 1 are a very small portion of the total market value and the unanticipated net flows will be an even smaller portion. The markets that we examine are all in developed countries where there should be adequate liquidity for absorbing such small amounts. Second, if there were temporary price pressure impacts from unanticipated flows, then there should be immediate reversals. However, the effects remain positive as H increases.

A second potential alternative explanation is that idiosyncratic shocks are being recorded as trades due to private information. In principle, these shocks need not correspond to news about future asset payoffs. Instead, there might be shocks to individual wealth or preferences.[39] However, any such candidate shock affects the interpretation of our results only if the average shock (i) has persistent price impact *and* (ii) is not observable. Condition (i) arises because our measures help predict returns.[40] It suggests that our results are not driven by simple preference shocks that are iid over time (though correlated in the cross section), for example generated by a group of "noise traders" exerting "price pressure". Condition (ii) arises because we identify the average trade by US investors due to private information that is orthogonal to public signals and, under the conservative measure, orthogonal also to current returns. It suggests that we are not simply picking up an average of shocks to individual wealth. Such an average should be closely related to returns, and would be counted as 'public' by our conservative measure.[41]

The third, and potentially most serious alternative explanation, is that we are counting trades due to public information as trades due to private information. As mentioned above, this is a potential problem with our broad measures of private information (where instruments have beginning-of-month values), but it is not a concern with our conservative measures of public information (where instruments have end-of-month values). Note that our conservative measures of private information are orthogonal to the realized return in the foreign asset market during the current month and in the previous three months. As mentioned above, using realized returns captures the impact of any public information variable that we have omitted from our flow regressions while also diminishing the private information effects. The fact that our conservative measures still have forecasting ability for future returns shows that our findings are not due to omitted public information.

Our analysis is also robust to omitted factors in the latent-factor model of returns.

[39] Indeed, we know that some trades by *some* agents have to be due to such shocks in order for *any* trades due to private signals to be valuable; i.e. signals are valuable only if prices do not reveal them. In the theoretical literature, this also requires that there is another unobservable factor, such as noise traders or supply shocks, driving prices. Of course, this factor need not affect trades by US investors.

[40] Of course, our results could be explained by serially correlated noise trader demands. But at some level, any pattern can be rationalized by a well-chosen noise distribution.

[41] The fact that our two measures are so similar thus confirms the finding of Bohn and Tesar (1996b) that portfolio rebalancing is not an important determinant of US investors' net trades.

If the missing factor is a function of the public information variables that we have used, then the broad measures of private information are orthogonal to the missing factor and our results go through. If the missing factor is a function of omitted public information variables, then the use of realized returns in the conservative measures of private information captures this effect.

Finally, a potential criticism is that the model of expected equity flows does not impose enough structure to separate private information from portfolio inventory shocks or other microstructure type shocks (e.g. smoothing of the price process by the specialist, stale quotes, etc.). These microstructure effects are important at very short horizons of one day or less whereas our data is monthly. In addition, theory suggests a demanding test for measures of private information. The key insight is that private information partially revealed through trades causes a persistent impact on a security's price. Thus, a shock to private information is different from other shocks by its persistent price impact. This is what Tables 6, 7, and 9 above showed for both the broad and conservative measures of private information.

4.5 Impact on Return Variation

We next construct variance-ratio measures of the importance of private information on realized return variation. The ratios are presented in Table 10. The denominator in each of these measures is the standard deviation of realized excess equity return in each of the countries. The numerator in the total (global) part of the table is the absolute value of the estimated γ_H ($\widetilde{\gamma}$) coefficients times the standard deviation of the corresponding private information variables. Again, both measures of private information are constructed using instrument set C (lagged values of the flows, plus global and local instruments). These ratios estimate the impact of a one standard deviation shock to the private information of US investors on international equity returns.

The table shows that the release of private information has a substantial effect on realized return variation over short holding periods. A one standard deviation shock to the total broad measure of private information of US investors results in an increase in contemporaneous realized equity returns ranging from 0.1 per cent for the UK up to 33.7 per cent for Japan. Over longer holding periods, the private information continues to have an impact. For a three month holding period, total private information accounts for 0.4 per cent of Swiss realized returns up to 17.9 per cent of realized Japanese equity returns. The impact of the conservative measures of private information are naturally smaller.

The table also indicates that the global component of the private information is capturing a considerable amount of the variation in realized returns. The amount of realized return variation explained by a one standard deviation shock to the global factor ranges from 5.2 per cent for Italy up to 24.3 per cent for Canada. Importantly, the global factor represents 24 per cent of realized US equity return variation. Over longer horizons, the impact is reduced but remains large for many countries.

5 Conclusions

Despite its increasing theoretical importance in accounting for several features of international equity flows and returns, there is little empirical evidence about the size and sources of private information and many important questions remain unanswered. This paper addresses some of these questions. We first formulate a portfolio choice problem where investors receive both public and private signals and identify trades due to private information. In order to obtain measures of the private information of US investors, we estimate an empirical model of international equity flows. The models of gross purchases and sales of foreign equities by US investors work extremely well with an average reported $\bar{R}^2$ of approximately 60 per cent across our sample of eight countries. We show that modeling gross flows results in a better model of expected net flows than modeling net purchases directly.

We find that our measures of private information have a significant impact on the cross section of international equity returns. Although the trades of US investors constitute a very small amount of the total market capitalization of the foreign country, these unexpected flows are able to predict returns. We show that our results are robust to potential "noise trading" by US investors, trading on omitted public information, and other potential model misspecifications.

Based on the model, we also suggest a way of identifying global factors in private information. We show that there is a factor structure in our measures of the total private information of US investors. The main factor, or global component, accounts for approximately 55 per cent of the variation in total private information and also has a significant joint impact on the cross section of international equity returns. Measured private information, total and global, is also large relative to realized return variation; a one standard error shock to the broad measure of total private information accounts for an average of 13.2 per cent of realized return variation across our eight foreign countries.

Our results suggest that US investors are likely to play a special role in international markets. Rather than suffering from an information disadvantage because of insufficient local information, they may enjoy an information advantage because of superior global information. This information advantage appears to be long lived and suggests potential trading strategies where (sophisticated) US investors outperform the average non-US investor participating in the foreign equity market. More analysis is required to see whether this is possible. Our analysis simply shows that US investors have significant private information about the returns on the indexes of these countries. We do not analyze how to take an optimal position in the cross section of firms in these markets.

We have chosen to work only with developed country data to avoid complications with transition periods associated with emerging stock markets. However, work by Morck et al. (2000) shows that there is greater price co-movement in these markets as opposed to developed markets. This suggests that our findings may be even stronger in these markets.

Finally, our results must be incorporated into any examination of the home bias puzzle. It is clear from our results that there is a group of US (institutional) investors who have significant private information about foreign equity markets and that a substantial portion of this information results from a global factor. The results in Tkac (2001) cited above show that they can outperform appropriate benchmarks. This raises two questions. The first is why do other US investors not hire these institutions to allocate their portfolios in international equity markets? One potential explanation is the existence of agency problems between the client and the fund manager, yet this seems to be a less than satisfactory explanation as these problems would also exist in a purely domestic setting. For example, a US investor choosing among (domestic) US mutual funds faces the problem of deciding whether the fund managers have significant ability, part of which will include private information. Why is that investor's choice any different with respect to choosing an international fund manager? In both cases, the fund managers will have private information that the investor may never be able to obtain.

The second question raised by our analysis is whether US investors have an absolute information advantage across all equity markets through their better knowledge of the global factor in private information. This finding is also not inconsistent with the home bias puzzle, as it is likely that US investors have a 'comparative information advantage' in their own equity market. More research is needed on the composition of private information for investors around the world.

A Derivation of Some Results

In this appendix we solve the individual investor's optimization problem presented in the main text. For simplicity we drop the superscript i that indexes the individual.

We start by guessing that the investor's lifetime utility as a function of beginning-of-period wealth and the information vector is given by:

$$V(w_t, I_t) = -\exp\left[-\kappa - \rho\frac{R^f - 1}{R^f} w_t - u' I_t - \frac{1}{2} I_t' U I_t\right], \tag{13}$$

with κ, u, and U being functions of the underlying model parameters. We solve for κ, u, and U below. The parameters u and U capture the impact on the investor's lifetime utility from changing state variables I_t^i. This is because the investor, being risk averse, cares not only about wealth fluctuations, but also changes in the payoff-relevant information vector.

Let M^{RI} denote a matrix of zeros and ones that extracts returns from the vector of payoff-relevant information; that is, $R_{t+1} = M^{RI} I_{t+1}^i$. Similarly, let M^{yI} be a matrix of zeros and ones that extracts non-asset income from the vector I; i.e., $y_{t+1}^i = M^{yI} I_{t+1}^i$. Also, define $\alpha = \rho\left(R^f - 1\right)/R^f$.

Using (2) and (13) we can evaluate the expected value of future lifetime utility following a consumption and investment strategy today (c_t, θ_t) and the optimal plan from tomorrow onwards:

$$\begin{aligned} E\left(V\left(w_{t+1}, I_{t+1}\right)|I_t\right) &= -E\left(\exp^{-\kappa - \alpha w_{t+1} - u' I_{t+1} - \frac{1}{2} I_{t+1}' U I_{t+1}} |I_t\right) \\ &= -\exp^{-\kappa - \rho(R^f - 1)(w_t - c_t)} E\left(\exp^{-\bar{u}_t I_{t+1} - \frac{1}{2} I_{t+1}' U I_{t+1}} |I_t\right), \end{aligned}$$

where $\bar{u}_t = \alpha[\theta_t' M^{RI} + M^{yI}] + u'$. Letting $d = |\Sigma_{II}|^{-\frac{1}{2}} / \left|\tilde{\Sigma}_{II}\right|^{-\frac{1}{2}}$ and $\tilde{\Sigma}_{II} = \left(\Sigma_{II}^{-1} + U\right)^{-1}$, we can write

$$\begin{aligned} \ln[-E\left(V\left(w_{t+1}, I_{t+1}\right)|I_t\right)/d] &= -\kappa - \rho(R^f - 1)\left(w_t - c_t\right) - \left[\bar{u}_t + \frac{1}{2} E\left(I_{t+1}|I_t\right)' U\right] E\left(I_{t+1}|I_t\right) \\ &\quad + \frac{1}{2}\left[\bar{u}_t + E\left(I_{t+1}|I_t\right)' U\right] \tilde{\Sigma}_{II} \left[\bar{u}_t + E\left(I_{t+1}|I_t\right)' U\right]'. \end{aligned}$$

We now turn to finding the solution to the investor's optimization problem:

$$V(w_t, I_t) = \max_{c_t, \theta_t}\left[-\exp^{-\rho c_t} + \delta E\left(V\left(w_{t+1}, I_{t+1}\right)|I_t\right)\right]. \tag{14}$$

The first order conditions to this problem are:

$$\begin{aligned}
\rho \exp^{-\rho c_t} &= \delta\rho(R^f-1)d\exp^{\pi(I_t)} \\
\alpha M^{RI}\tilde{\Sigma}_{II}M^{RI\prime}\theta_t &= M^{RI}\left(I-\tilde{\Sigma}_{II}U\right)\Phi_1 I_t + M^{RI}\Phi_0 - M^{RI}\tilde{\Sigma}_{II}\left(\alpha M^{yI\prime}+u+U\Phi_0\right).
\end{aligned}$$

These first order conditions can be solved to obtain:

$$\begin{aligned}
c_t &= \bar{c}_t + \rho\frac{R^f-1}{R^f}w_t \\
\theta_t &= \bar{\theta} + \Theta I_t,
\end{aligned} \tag{15}$$

with the matrix Θ, vector $\bar{\theta}$, and parameter $\bar{c}_t$ given by:

$$\begin{aligned}
\Theta &= \frac{R^f}{\rho(R^f-1)}\left[M^{RI}\tilde{\Sigma}_{II}M^{RI\prime}\right]^{-1}\left[M^{RI}\left(I-\tilde{\Sigma}_{II}U\right)\Phi_1\right], \\
\bar{\theta} &= \frac{R^f}{\rho(R^f-1)}\left[M^{RI}\tilde{\Sigma}_{II}M^{RI\prime}\right]^{-1}\left[M^{RI}\Phi_0 - M^{RI}\tilde{\Sigma}_{II}\left(\rho\frac{R^f-1}{R^f}M^{yI\prime}+u+U\Phi_0\right)\right],
\end{aligned}$$

and

$$\bar{c}_t = \frac{1}{\rho R^f}\left[-\log\left(\delta d\left(R^f-1\right)\right) - \pi\left(I_t\right)\right].$$

In $\bar{c}_t$, $\pi\left(I_t\right)$ is given by

$$\begin{aligned}
\pi\left(I_t\right) &= -\kappa - \frac{1}{2}\alpha^2\left(\bar{\theta}+\Theta I_t\right)' M^{RI}\tilde{\Sigma}_{II}M^{RI\prime}\left(\bar{\theta}+\Theta I_t\right) \\
&\quad - \left(\alpha M^{yI} + u' + \frac{1}{2}E\left(I_{t+1}|I_t\right)' U\right)E\left(I_{t+1}|I_t\right) \\
&\quad + \frac{1}{2}\left[\alpha M^{yI} + u' + E\left(I_{t+1}|I_t\right)' U\right]\tilde{\Sigma}_{II}\left[\alpha M^{yI} + u' + E\left(I_{t+1}|I_t\right)' U\right]'.
\end{aligned}$$

The parameter $\bar{c}_t$ is time-varying because consumption responds to movements in the information vector.

A convenient way to express the asset demands is to use a modified conditional distribution for tomorrow's information. Define

$$\tilde{I}^i_{t+1}|I^i_t \sim N\left(\Phi_0 + \Phi_1 I^i_t, \tilde{\Sigma}_{II}\right). \tag{16}$$

Then the optimal asset allocation (15) can be represented as

$$\begin{aligned}
\theta^i_t &= \frac{R^f}{\rho\left(R^f-1\right)}(M^{RI}\tilde{\Sigma}_{II}M^{RI\prime})^{-1}\left[E\left(M^{RI}\tilde{I}^i_{t+1}|I^i_t\right)\right. \\
&\quad \left. + cov\left([\rho\frac{R^f-1}{R^f}M^{yI} + u' + E\left(\tilde{I}^i_{t+1}|I^i_t\right)' U]\tilde{I}^i_{t+1}, M^{RI}\tilde{I}^i_{t+1}|I^i_t\right)\right].
\end{aligned} \tag{17}$$

If the investor does not hedge fluctuations in any of the state variables ($U = 0$ and $u = 0$, so $\tilde{\Sigma}_{II} = \Sigma_{II}$, and the investor ignores fluctuations in labor income), the second term in (17) disappears while the first term reduces to the familiar expression for the portfolio demands of a myopic investor. If the investor is not myopic, the second term is present as the investor hedges against future changes in payoff relevant information. In our homoskedastic normal setup, both terms are *linear* in current information I_t^i. Finally, the term $M^{RI}\tilde{\Sigma}_{II}M^{RI\prime}$ is the modified conditional variance of returns.

Finally, we need to solve for the parameters κ, u, and U and show that they are time-invariant. Feeding the optimal consumption and asset allocation decisions back into the Belman equation (14), we see that the parameters u and U solve a set of simultaneous equations:

$$U = \alpha^2 \Theta' M^{RI} \tilde{\Sigma}_{II} M^{RI\prime} \Theta + \Phi_1' U \Phi_1 - \Phi_1' U \tilde{\Sigma}_{II} U \Phi_1,$$

$$\begin{aligned} u' &= \left[I - \Phi_1 + \tilde{\Sigma}_{II} U \Phi_1 \right]^{-1} \left[\alpha^2 \bar{\theta}' M^{RI} \tilde{\Sigma}_{II} M^{RI\prime} \Theta + \alpha M^{yI} \Phi_1 \right. \\ &\quad \left. - \left[\alpha M^{yI} \tilde{\Sigma}_{II} + \Phi_0' U \tilde{\Sigma}_{II} - \Phi_0' \right] U \Phi_1 \right], \end{aligned}$$

and that the constant κ is

$$\begin{aligned} \left(1 + R^f\right) \kappa &= -R^f \log\left(\rho/\alpha\right) - \log\left(\delta d \left(R^f - 1\right)\right) - \frac{1}{2} \alpha^2 \bar{\theta}' M^{RI} \tilde{\Sigma}_{II} M^{RI\prime} \bar{\theta} \\ &\quad - \left[\alpha M^{yI} + u' + \frac{1}{2} \Phi_0' U \right] \Phi_0 \\ &\quad + \frac{1}{2} \left[\alpha M^{yI} + u' + \Phi_0' U \right] \tilde{\Sigma}_{II} \left[\alpha M^{yI} + u' + \Phi_0' U \right]'. \end{aligned}$$

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Table 1a
Summary Statistics of US Investors' International Portfolio Equity Flows

The table shows summary statistics of US investors' gross purchases, gross sales, and net purchases of foreign equities from Germany, Japan, UK, France, Canada, Netherlands, Switzerland and Italy. All of the series are expressed in monthly per-cent continuously compounded form. The flows data have been normalized by dividing by the market value of the foreign equity index at the start of the month, both expressed in US dollars. The sample period is January, 1977 to September, 2000. The three estimated autocorrelations of the series are for a one, six and twelve month lag, respectively.

	Mean (%)	Standard Deviation (%)	Autocorrelation		
			ρ_1	ρ_6	ρ_{12}
US Investors' Gross Purchases of Foreign Equities					
Germany	0.178	0.146	0.856	0.746	0.671
Japan	0.135	0.129	0.930	0.857	0.803
UK	0.718	0.574	0.954	0.915	0.895
France	0.265	0.126	0.470	0.332	0.109
Canada	0.589	0.364	0.828	0.642	0.480
Netherlands	0.237	0.145	0.672	0.519	0.386
Switzerland	0.396	0.193	0.647	0.379	0.325
Italy	0.142	0.164	0.597	0.489	0.420
US Investors' Gross Sales of Foreign Equities					
Germany	0.168	0.144	0.907	0.799	0.736
Japan	0.121	0.115	0.949	0.917	0.857
UK	0.681	0.576	0.967	0.934	0.910
France	0.216	0.095	0.444	0.275	0.185
Canada	0.534	0.341	0.841	0.740	0.577
Netherlands	0.222	0.142	0.737	0.669	0.562
Switzerland	0.354	0.174	0.500	0.301	0.158
Italy	0.128	0.156	0.559	0.432	0.339
US Investors' Net Purchases of Foreign Equities					
Germany	0.011	0.062	0.313	0.099	0.118
Japan	0.014	0.048	0.571	0.060	-0.098
UK	0.037	0.101	0.436	0.222	0.232
France	0.048	0.118	0.265	0.106	-0.037
Canada	0.055	0.162	0.447	0.056	-0.022
Netherlands	0.016	0.103	0.353	0.028	-0.058
Switzerland	0.042	0.160	0.292	-0.025	-0.023
Italy	0.014	0.149	0.148	0.055	-0.020

Table 1b
Summary Statistics of Excess Equity Returns and Global Instruments

The table shows summary statistics of: (i) excess equity returns from the US, Germany, Japan, UK, France, Canada, Netherlands, Switzerland, and Italy; and (ii) the global instrument set: the US short-term interest rate (r_t^{US}), the US credit spread ($crsp_t^{US}$), the dividend yield on the global equity market (div_t^{all}), and the slope of the US term structure (sl_t^{US}). All of the series are expressed in monthly per cent continuously compounded form. The sample period is January, 1977 to September, 2000. The three estimated autocorrelations of the series are for a one, six and twelve month lag, respectively.

	Mean (%)	Standard Deviation (%)	Autocorrelation ρ_1	ρ_6	ρ_{12}
Excess equity returns					
US	0.605	4.265	-0.009	-0.053	0.032
Germany	0.356	5.548	-0.043	0.024	-0.033
Japan	0.329	6.661	0.072	-0.029	0.037
UK	0.712	5.431	-0.056	-0.094	-0.118
France	0.719	6.526	0.030	-0.022	-0.088
Canada	0.354	5.243	0.031	0.026	-0.114
Netherlands	0.722	4.623	-0.071	-0.050	0.028
Switzerland	0.553	5.152	0.048	-0.075	-0.011
Italy	0.538	7.519	0.071	0.082	0.057
Global instruments					
r_t^{US}	0.660	0.295	0.961	0.837	0.742
$crsp_t^{US}$	0.091	0.039	0.967	0.833	0.396
div_t^{all}	0.225	0.087	0.994	0.962	0.916
sl_t^{US}	0.185	3.341	0.144	0.039	-0.021

Table 1c
Correlations of US Investors' Gross Purchases, Gross Sales, and Net Purchases of Foreign Equities with Foreign Excess Stock Returns

The table shows correlation coefficients of US investors' gross purchases, gross sales, and net purchases of foreign equities with the excess return on the stock market in that country. The flows data are presented in Table 1a while the excess returns are given in Table 1b.

	Gross Purchases	Gross Sales	Net Purchases
Excess return in:			
Germany	0.120	0.040	0.187
Japan	0.130	0.005	0.340
UK	0.003	-0.006	0.048
France	0.111	0.016	0.105
Canada	0.152	0.092	0.148
Netherlands	0.153	0.089	0.092
Switzerland	0.022	-0.012	0.039
Italy	0.141	0.052	0.101

Table 2a
Summary Statistics from OLS Regressions of US Investors' Gross Purchases of Foreign Equities on Different Information Sets

The table presents summary statistics from OLS regressions of gross purchases of foreign equities by US investors on four information sets: (A) three lagged values of the gross purchases; (B) lagged purchases plus the global instruments; (C) lagged purchases plus global and local instruments; and (D) lagged purchases plus lagged squared residuals from a vector autoregression of the US excess stock return, the foreign country excess stock return and the change in the exchange rate. The global instruments are detailed in Table 1b. The local instruments (all lagged three periods) include: the foreign country's stock return; the foreign country's dividend yield; the difference in interest rates between the foreign country and the US; and the change in the spot exchange rate of the country's currency against the US dollar. The $\overline{R}^2$ statistics are adjusted for degrees of freedom. The value of the chi-squared test statistic associated with the Wald test of the null hypothesis that the coefficients on the explanatory variables are jointly equal to zero is shown in the column (χ^2) along with its asymptotic marginal significance level (*P*-value). In some of the tests, the degrees of freedom are adjusted lower due to multicollinearity of the regressors.

	Instrument Set							
	(A) lagged flow		(B) lagged flow + global instruments		(C) lagged flow + global + local instruments		(D) lagged flow + squared residuals	
	$\overline{R}^2$	$\chi^2(3)$ for lagged flows *P*-value	$\overline{R}^2$	$\chi^2(12)$ for global variables *P*-value	$\overline{R}^2$	$\chi^2(12)$ for local variables *P*-value	$\overline{R}^2$	$\chi^2(3)$ for squared residuals *P*-value
Germany	0.759	1111.25 0.000	0.770	38.54 0.000	0.767	3.12 0.539	0.758	3.88 0.274
Japan	0.891	4771.24 0.000	0.889	7.62 0.746	0.896	3201.72 0.000	0.891	5.09 0.165
UK	0.933	310.01 0.000	0.933	141.06 0.000	0.935	140.38 0.000	0.934	9.53 0.022
France	0.314	147.08 0.000	0.352	62.50 0.000	0.353	20.52 0.058	0.314	8.14 0.043
Canada	0.715	1423.67 0.000	0.715	34.23 0.000	0.717	32.84 0.001	0.713	2.394 0.495
Netherlands	0.538	452.48 0.000	0.562	793.06 0.000	0.585	32.76 0.001	0.535	2.23 0.525
Switzerland	0.476	195.48 0.000	0.504	178.00 0.000	0.517	442.82 0.000	0.486	6.10 0.106
Italy	0.446	305.09 0.000	0.471	16.84 0.051	0.467	22.39 0.033	0.445	5.10 0.164

Table 2b
Summary Statistics from OLS Regressions of US Investors' Gross Sales of Foreign Equities on Different Information Sets

The table presents summary statistics from OLS regressions of gross sales of foreign equities by US investors on four information sets: (A) three lagged values of the gross sales; (B) lagged sales plus the global instruments; (C) lagged sales plus global and local instruments; and (D) lagged sales plus lagged squared residuals from a vector autoregression of the US excess stock return, the foreign country excess stock return and the change in the exchange rate. The global instruments are detailed in Table 1b while the local instruments are detailed in Table 2a. The $\overline{R}^2$ statistics are adjusted for degrees of freedom. The value of the chi-squared test statistic associated with the Wald test of the null hypothesis that the coefficients on the explanatory variables are jointly equal to zero is shown in the column (χ^2) along with its asymptotic marginal significance level (*P*-value). In some of the tests, the degrees of freedom are adjusted lower due to multicollinearity of the regressors.

	Instrument Set							
	(A) lagged flow		(B) lagged flow + global instruments		(C) lagged flow + global + local instruments		(D) lagged flow + squared residuals	
	$\overline{R}^2$	$\chi^2(3)$ lagged flow *P*-value	$\overline{R}^2$	$\chi^2(12)$ for global variables *P*-value	$\overline{R}^2$	$\chi^2(12)$ for local variables *P*-value	$\overline{R}^2$	$\chi^2(3)$ for squared residuals *P*-value
Germany	0.844	1326.14 0.000	0.845	102.54 0.000	0.848	12.72 0.122	0.844	3.54 0.316
Japan	0.921	6976.32 0.000	0.919	18.45 0.072	0.922	169.81 0.000	0.921	5.12 0.163
UK	0.954	123.83 0.000	0.954	112.62 0.000	0.954	21.10 0.012	0.955	5.88 0.118
France	0.242	98.69 0.000	0.295	63.69 0.000	0.276	86.81 0.000	0.236	0.77 0.857
Canada	0.755	1091.38 0.000	0.749	228.32 0.000	0.748	8.73 0.726	0.753	1.16 0.763
Netherlands	0.638	723.24 0.000	0.658	30.33 0.002	0.674	31.37 0.002	0.636	5.00 0.172
Switzerland	0.319	67.77 0.000	0.367	43.32 0.000	0.383	31.43 0.002	0.314	1.66 0.646
Italy	0.398	64.44 0.000	0.413	63.14 0.000	0.411	24.10 0.020	0.395	4.54 0.209

Table 2c
Summary Statistics from OLS Regressions of US Investor's Net Purchases of Foreign Equities on Different Information Sets

The table presents summary statistics from OLS regressions of net purchases of foreign equities by US investors on four information sets: (A) three lagged values of the net purchases; (B) lagged net purchases plus the global instruments; (C) lagged net purchases plus global and local instruments; and (D) lagged net purchases plus lagged squared residuals from a vector autoregression of the US excess stock return, the foreign country excess stock return and the change in the exchange rate. The global instruments are detailed in Table 1b while the local instruments are detailed in Table 2a. The $\overline{R}^2$ statistics are adjusted for degrees of freedom. The value of the chi-squared test statistic associated with the Wald test of the null hypothesis that the coefficients on the explanatory variables are jointly equal to zero is shown in the column (χ^2) along with its asymptotic marginal significance level (*P*-value). In some of the tests, the degrees of freedom are adjusted lower due to multicollinearity of the regressors.

	Instrument Set							
	(A) lagged flow		(B) lagged flow + global instruments		(C) lagged flow + global + local instruments		(D) lagged flow + squared residuals	
	$\overline{R}^2$	$\chi^2(3)$ lagged flow *P*-value	$\overline{R}^2$	$\chi^2(12)$ for global variables *P*-value	$\overline{R}^2$	$\chi^2(12)$ for local variables *P*-value	$\overline{R}^2$	$\chi^2(3)$ for squared residuals *P*-value
Germany	0.116	55.12	0.103	18.05	0.094	6.27	0.135	21.22
		0.000		0.114		0.712		0.000
Japan	0.338	915.95	0.322	12.36	0.348	77.95	0.339	3.83
		0.000		0.417		0.000		0.280
UK	0.227	73.47	0.226	31.13	0.291	871.44	0.228	13.77
		0.000		0.003		0.000		0.003
France	0.115	21.14	0.153	45.65	0.143	16.38	0.124	8.48
		0.000		0.000		0.128		0.037
Canada	0.221	73.09	0.228	75.52	0.249	23.12	0.219	3.70
		0.000		0.000		0.027		0.296
Netherlands	0.166	62.94	0.191	65.65	0.246	35.01	0.170	21.09
		0.000		0.000		0.000		0.000
Switzerland	0.097	12.78	0.086	14.23	0.118	259.42	0.117	6.49
		0.005		0.286		0.000		0.090
Italy	0.021	9.66	0.001	28.46	0.001	17.73	0.014	4.84
		0.021		0.003		0.088		0.183

Table 3
Alternative Methods of Modeling Expected Net Flows

The table shows two statistics to compare modeling expected net flows either (i) directly in an OLS regression or (ii) indirectly via two separate regressions of the gross purchases and sales. The first number is the ratio of the variance of the expected net flow series constructed by two methods. In the numerator, the expected net flow is obtained from a regression of the net flows on the given instrument set, as in Table 2c. In the denominator, the expected portion of the gross flows regressions from Tables 2a and 2b are used to construct an implied expected net flow. A number below 1.00 indicates that we obtain more explanatory power for net flows by using the expected gross flows to construct an expected net flow rather than by using the net flow in a regression directly. The second number (in parentheses) is the asymptotic marginal significance level (*P*-value) of a test of the two methods. The net flow is regressed on the given information set with the lagged gross purchases and sales variables entering separately. The test is that the coefficients on the lagged purchases are equal (and opposite in sign) to the coefficients on the lagged gross sales. Each test statistic is χ^2 distributed with three degrees of freedom.

	Instrument Set			
	(A) lagged flow	(B) lagged flow + global instruments	(C) lagged flow + global + local instruments	(D) lagged flow + squared residuals
Country	variance ratio (*P*-value)	variance ratio (*P*-value)	variance ratio (*P*-value)	variance ratio (*P*-value)
Germany	0.235 (0.003)	0.336 (0.143)	0.417 (0.055)	0.276 (0.014)
Japan	0.487 (0.048)	0.536 (0.081)	0.670 (0.048)	0.501 (0.119)
UK	0.367 (0.000)	0.412 (0.000)	0.526 (0.000)	0.388 (0.000)
France	0.507 (0.998)	0.739 (0.996)	0.802 (0.999)	0.540 (0.956)
Canada	0.416 (0.865)	0.501 (0.956)	0.624 (0.856)	0.426 (0.881)
Netherlands	0.379 (0.646)	0.896 (0.909)	0.992 (0.997)	0.391 (0.584)
Switzerland	0.307 (0.000)	0.542 (0.000)	0.762 (0.000)	0.360 (0.001)
Italy	0.113 (0.326)	0.296 (0.098)	0.472 (0.056)	0.130 (0.205)

Table 4
Summary Statistics of Regressions of International Excess Equity Returns on Global and Local Instruments

The table shows summary statistics from OLS regressions of the excess equity returns on US and foreign stocks on two instrument sets: (i) the set of global instruments shown in Table 1b; and (ii) the global and local instruments. The $\overline{R}^2$ statistics are adjusted for degrees of freedom. The value of the chi-squared test statistic associated with the Wald test of the null hypothesis that the coefficients on the explanatory variables are jointly equal to zero is shown in the column (χ^2) along with its asymptotic marginal significance level (*P*-value).

	(i) Global instruments		(ii) Global plus local instruments	
Country	$\overline{R}^2$	χ^2 (4) *P*-value	$\overline{R}^2$	χ^2 (4) for local *P*-value
US	0.019	9.301 0.054	0.018	0.815 0.665
Germany	0.038	14.654 0.005	0.046	5.868 0.209
Japan	0.032	13.455 0.009	0.030	4.873 0.301
UK	0.016	6.451 0.168	0.037	8.104 0.088
France	0.028	7.044 0.134	0.018	0.649 0.957
Canada	0.022	6.162 0.187	0.016	3.408 0.492
Netherlands	0.041	10.603 0.031	0.052	4.919 0.296
Switzerland	0.065	18.727 0.001	0.055	0.657 0.957
Italy	-0.003	3.128 0.537	-0.013	1.246 0.870

Table 5a
Coefficients on Global Instruments in the One-Factor Latent-Variable Model of International Stock Returns

The table shows the α_H coefficients on the global instruments in the latent-variable model of international stock returns. The global instruments are a constant, the US short-term interest rate (r_{t-1}^{US}), the US credit spread ($crsp_{t-1}^{US}$), the dividend yield on the global equity market (div_{t-1}^{all}), and the slope of the US term structure (sl_{t-1}^{US}). The model is estimated separately by Generalized Method of Moments for each holding period of H months. The beta coefficient on the US stock return is normalized to 1.00 for identification. The small-sample adjusted version of the Newey-West standard errors (s.e.) are calculated assuming an overlap of $H+3$ terms in the error process. The small-sample adjustment follows the procedure outlined in Ferson and Foerster (1994). The small-sample marginal significance levels (P-values) are shown below the estimates.

	α_H coefficient in latent variable model				
Holding period	constant (s.e.) P-value	r_{t-1}^{US} (s.e.) P-value	$crsp_{t-1}^{US}$ (s.e.) P-value	div_{t-1}^{all} (s.e.) P-value	sl_{t-1}^{US} (s.e.) P-value
$H = 0$	0.004	-3.403	10.002	5.826	0.178
	(0.004)	(0.836)	(5.053)	(2.422)	(0.060)
	0.365	0.000	0.049	0.017	0.003
$H = 1$	0.007	-6.508	18.477	11.637	0.156
	(0.005)	(1.374)	(6.127)	(3.754)	(0.052)
	0.174	0.000	0.003	0.002	0.003
$H = 2$	0.001	-9.920	42.638	18.477	0.206
	(0.008)	(1.581)	(9.187)	(4.826)	(0.068)
	0.903	0.000	0.000	0.000	0.003
$H = 3$	0.001	-11.120	41.899	21.918	0.132
	(0.007)	(2.032)	(8.954)	(5.853)	(0.058)
	0.888	0.000	0.000	0.000	0.024

Table 5b
Coefficients on the Implied Global Risk Premium in the One-Factor Latent-Variable Model of International Stock Returns

The table shows the β_H coefficients on the implied global risk premium from the latent-variable model of international stock returns. The implied global risk premium is the linear combination of the instruments given in Table 5a. The beta coefficient on the US stock return is normalized to 1.00 for identification. The model is estimated separately for each holding period of *H* months by Generalized Method of Moments. The small-sample adjusted version of the Newey-West standard errors (s.e.) are calculated assuming an overlap of *H+3* terms in the error process. The small-sample marginal significance levels (*P*-values) are shown below the estimates. The small-sample adjustment follows the procedure outlined in Ferson and Foerster (1994).

	Holding period			
	H = 0	*H = 1*	*H = 2*	*H = 3*
Country	β (s.e.) *P*-value	β (s.e.) *P*-value	β (s.e.) *P*-value	β (s.e.) *P*-value
US	1.000	1.000	1.000	1.000
Germany	1.291	1.392	0.829	1.167
	(0.326)	(0.347)	(0.238)	(0.334)
	0.000	0.000	0.001	0.001
Japan	1.882	2.289	2.523	3.046
	(0.539)	(0.542)	(0.416)	(0.578)
	0.001	0.000	0.000	0.000
UK	0.997	1.116	0.861	0.791
	(0.267)	(0.262)	(0.176)	(0.193)
	0.000	0.000	0.000	0.000
France	1.183	1.518	1.361	1.569
	(0.407)	(0.359)	(0.261)	(0.292)
	0.004	0.000	0.000	0.000
Canada	1.306	1.010	1.106	1.107
	(0.202)	(0.203)	(0.126)	(0.153)
	0.000	0.000	0.000	0.000
Netherlands	1.418	1.637	1.302	1.617
	(0.289)	(0.280)	(0.158)	(0.229)
	0.000	0.000	0.000	0.000
Switzerland	1.545	1.947	1.405	1.734
	(0.360)	(0.417)	(0.257)	(0.347)
	0.000	0.000	0.000	0.000
Italy	1.175	1.016	1.230	1.500
	(0.424)	(0.476)	(0.355)	(0.396)
	0.006	0.034	0.001	0.000

Table 5c
Summary Statistics of the One-Factor Latent-Variable Model of International Stock Returns

The top part of the table presents variance-ratio measures of the statistical fit of the latent-variable model. The ratio shows how the latent-variable model without any private information captures the expected return variation in the data. The numerator is the variance of the expected return from the latent-variable model while the denominator is the variance of the expected return from an OLS regression of the return on the global instruments. The bottom part of the table presents the value of the *J*-statistic used to test the over-identifying restrictions of the model. The statistics are distributed as $\chi^2(32)$ and are presented along with their asymptotic marginal significance levels (*P*-value).

	Holding period			
Country	*H = 0*	*H = 1*	*H = 2*	*H = 3*
Variance ratios				
US	1.541	1.062	1.759	1.356
Germany	0.971	0.772	0.590	0.772
Japan	1.643	1.790	2.599	2.350
UK	0.940	0.576	0.628	0.472
France	0.712	0.607	0.629	0.469
Canada	1.587	0.688	1.346	1.142
Netherlands	1.681	1.264	1.350	1.413
Switzerland	1.083	1.007	0.803	0.848
Italy	2.131	0.681	1.412	1.098
J-statistic model test				
$\chi^2(32)$	24.111	27.717	27.639	27.224
P-value	0.840	0.683	0.687	0.707

Table 6
Joint Significance Tests of the Effects of the Broad Measures of Private Information on Equity Returns across all Foreign Countries

The table shows the values of the chi-squared test statistics associated with the Wald tests of the null hypothesis that all of the γ_H coefficients on the broad private information measures are jointly equal to zero. In the first part of the table, the tests are for the joint significance of the total measures of private information of US investors across the equity returns in the eight foreign country (Germany, Japan, UK, France, Canada, Netherlands, Switzerland and Italy) that are presented in Table 7. The statistics are $\chi^2(8)$ distributed. In the second part of the table, the tests are for the joint significance of the global factor in the broad measures of private information of US investors across the equity returns of the eight foreign countries plus the US that are presented in Table 9. The statistics are $\chi^2(9)$ distributed. Both statistics are obtained from a Generalized Method of Moments estimation of the model and presented with their asymptotic marginal significance levels (*P*-values).

	Instrument Set			
	(A) lagged flow	(B) lagged flow + global instruments	(C) lagged flow + global + local instruments	(D) lagged flow + squared residuals
	Total private information ($\chi^2(8)$ and *P*-value)			
H = 0	445.522 0.000	504.791 0.000	576.010 0.000	440.233 0.000
H = 1	299.440 0.000	312.646 0.000	324.634 0.000	305.514 0.000
H = 2	332.473 0.000	517.148 0.000	539.216 0.000	287.694 0.000
H = 3	420.533 0.000	458.945 0.000	442.890 0.000	375.880 0.000
	Global factor in private information ($\chi^2(9)$ and *P*-value)			
H = 0	64.109 0.000	58.364 0.000	62.052 0.000	55.245 0.000
H = 1	37.083 0.000	30.888 0.000	29.634 0.001	27.226 0.001
H = 2	30.570 0.000	33.488 0.000	39.173 0.000	17.655 0.039
H = 3	20.073 0.017	30.079 0.000	47.492 0.000	15.609 0.076

Table 7

Coefficients on Measures of Total Private Information of US Investors in International Equity Markets

In the left panel, the table shows the γ_H coefficients on the broad measures of total private information of US investors in the cross section of international stock returns. In the right panel, the table shows the $\tilde{\gamma}$ coefficients on the conservative measures of total private information of US investors in the cross section of international stock returns. In both panels, the coefficients are estimated separately for each holding period of H months but jointly across the equity returns of all of the foreign countries as in equation (10) by Generalized Method of Moments. The α_H and β_H coefficients are not reported as they are similar to those presented in Tables 5a and 5b. The small-sample adjusted version of the Newey-West standard errors (s.e.) are calculated assuming an overlap of $H+3$ terms in the error process. The small-sample marginal significance levels (P-values) are shown below the estimates. The small-sample adjustment follows the procedure outlined in Ferson and Foerster (1994). The values of the chi-squared test statistics associated with the Wald tests of the null hypothesis that all of the γ coefficients on the private information measures are jointly equal to zero are shown at the bottom of the table.

	Holding period for broad measure				Conservative measure
	$H = 0$	$H = 1$	$H = 2$	$H = 3$	
Country	γ_0 (s.e.) P-value	γ_1 (s.e.) P-value	γ_2 (s.e.) P-value	γ_3 (s.e.) P-value	$\tilde{\gamma}$ (s.e.) P-value
US	-	-	-	-	
Germany	12.776	17.266	21.430	16.194	1.828
	(2.105)	(2.679)	(3.174)	(3.079)	2.089
	0.000	0.000	0.000	0.000	0.382
Japan	61.478	70.207	65.400	69.381	25.082
	(5.790)	(7.266)	(8.118)	(10.721)	6.546
	0.000	0.000	0.000	0.000	0.000
UK	0.068	1.813	14.921	3.945	4.643
	(1.698)	(1.959)	(1.980)	(2.257)	1.851
	0.968	0.356	0.000	0.082	0.013
France	8.521	5.629	12.498	16.126	0.401
	(1.517)	(1.390)	(1.779)	(1.783)	1.193
	0.000	0.000	0.000	0.000	0.737
Canada	9.381	5.847	6.788	5.501	-1.390
	(1.327)	(1.306)	(1.075)	(1.049)	0.984
	0.000	0.000	0.000	0.000	0.159
Netherlands	7.539	1.425	9.584	8.665	-2.244
	(0.938)	(1.060)	(1.135)	(1.210)	0.969
	0.000	0.180	0.000	0.000	0.021
Switzerland	1.210	2.833	2.476	-0.328	4.531
	(0.676)	(0.966)	(0.921)	(0.899)	1.144
	0.075	0.004	0.008	0.715	0.000
Italy	1.651	10.624	6.121	3.839	0.338
	(1.055)	(2.900)	(1.610)	(1.885)	1.466
	0.119	0.000	0.000	0.043	0.818
Model test statistic					
$\chi^2(8)$	576.010	324.634	539.216	442.890	84.109
P-value	0.000	0.000	0.000	0.000	0.000

Table 8
Factor Analysis of Broad Measures of Private Information in International Portfolio Equity Flows

The table presents results of a factor analysis on the residuals from the gross flows regressions using beginning-of-month values of instrument set C (lagged flows, plus global and local instruments) shown in Tables 2a and 2b. The first column presents the cumulative variance of residuals explained by the first three factors obtained from an iterated principal factor analysis. The second column presents a chi-squared test statistic that the covariance matrix does not display a factor structure against an alternative one factor representation. The third column presents a chi-squared test statistic that the covariance matrix contains more than one factor. The latter two tests are obtained from a maximum-likelihood analysis and are presented along with their marginal significance levels (*P*-values).

Factor	Cumulative Variance (%)	Test of 0 factors against 1 factor $\chi^2(8)$ *P*-value	Test of >1 factor against 1 factor $\chi^2(20)$ *P*-value
Gross Purchases			
1	0.569	158.98	44.65
2	0.748	0.000	0.001
3	0.908		
Gross Sales			
1	0.533	139.56	31.00
2	0.683	0.000	0.055
3	0.814		

Table 9

Coefficients on Global Factor in Private Information in International Equity Markets

In the first panel, the table shows the γ_H coefficients on the broad measures of the global factor in private information of US investors in the cross section of international stock returns. In the second panel, the table shows the $\tilde{\gamma}$ coefficients on the conservative measures of the global factor in private information of US investors in the cross section of international stock returns. The estimation method is the same as that detailed in Table 7. The small-sample adjusted version of the Newey-West standard errors (s.e.) are calculated assuming an overlap of *H+3* terms in the error process. The small-sample marginal significance levels (*P*-values) are shown below the estimates. The small-sample adjustment follows the procedure outlined in Ferson and Foerster (1994). The values of the chi-squared test statistics associated with the Wald tests of the null hypothesis that all of the γ coefficients on the private information measures are jointly equal to zero are shown at the bottom of the table.

	Holding period for broad measure				Conservative measure
	H = 0	*H = 1*	*H = 2*	*H = 3*	
	γ_0	γ_1	γ_2	γ_3	$\tilde{\gamma}$
	(s.e.)	(s.e.)	(s.e.)	(s.e.)	(s.e.)
Country	*P*-value	*P*-value	*P*-value	*P*-value	*P*-value
US	15.829	19.753	22.185	19.219	4.947
	(4.179)	(6.114)	(6.154)	(5.908)	4.249
	0.000	0.001	0.000	0.001	0.245
Germany	7.564	-0.151	6.720	5.756	-6.024
	(5.251)	(6.698)	(7.827)	(7.740)	6.519
	0.151	0.982	0.391	0.458	0.356
Japan	6.155	9.566	-0.534	-9.881	2.616
	(7.106)	(9.755)	(11.305)	(11.374)	7.428
	0.387	0.328	0.962	0.386	0.725
UK	14.030	16.843	21.770	13.763	2.183
	(5.219)	(6.159)	(5.792)	(7.365)	5.329
	0.008	0.007	0.000	0.063	0.682
France	11.695	4.601	6.487	0.265	-6.916
	(5.439)	(6.076)	(7.839)	(8.206)	5.968
	0.032	0.450	0.409	0.974	0.248
Canada	19.698	18.240	14.660	11.148	1.157
	(4.395)	(6.388)	(6.981)	(7.806)	5.771
	0.000	0.005	0.037	0.155	0.841
Netherlands	16.096	9.754	12.419	8.419	-6.176
	(4.355)	(5.081)	(5.740)	(6.348)	4.688
	0.000	0.056	0.031	0.186	0.189
Switzerland	14.947	7.039	9.421	7.309	-8.558
	(4.712)	(6.980)	(7.936)	(9.479)	5.461
	0.002	0.314	0.236	0.441	0.118
Italy	6.067	2.019	7.488	17.449	-6.985
	(7.562)	(8.366)	(10.336)	(11.443)	6.836
	0.423	0.809	0.469	0.127	0.308
Model test statistic					
$\chi^2(9)$	62.052	29.634	39.173	47.492	34.684
P-value	0.000	0.001	0.000	0.000	0.000

Table 10
Variance-Ratio Measures of Private Information in International Equity Markets

The numerator is the (absolute value of) the γ coefficients from Tables 7 and 9 multiplied by the standard deviation of the corresponding measure of private information. The denominator is the standard deviation of the realized excess stock return in the country.

	Holding period for broad measure				
Country	*H = 0*	*H = 1*	*H = 2*	*H = 3*	Conservative measure
Total private information					
US	-	-	-	-	-
Germany	0.134	0.131	0.133	0.085	0.019
Japan	0.337	0.263	0.199	0.179	0.125
UK	0.001	0.020	0.144	0.033	0.071
France	0.138	0.064	0.118	0.130	0.006
Canada	0.245	0.107	0.101	0.071	0.034
Netherlands	0.137	0.019	0.107	0.085	0.039
Switzerland	0.034	0.055	0.039	0.004	0.125
Italy	0.032	0.141	0.067	0.035	0.006
Global factor in private information					
US	0.240	0.212	0.198	0.148	0.076
Germany	0.088	0.001	0.046	0.033	0.071
Japan	0.060	0.063	0.003	0.044	0.026
UK	0.167	0.145	0.160	0.089	0.026
France	0.117	0.032	0.037	0.001	0.070
Canada	0.243	0.156	0.103	0.066	0.014
Netherlands	0.227	0.100	0.107	0.062	0.088
Switzerland	0.188	0.061	0.065	0.042	0.109
Italy	0.052	0.012	0.036	0.069	0.061