# Treasury Bond Illiquidity and Global Equity Returns

Ruslan Goyenko and Sergei Sarkissian *

April 20, 2011

* The authors are from the Faculty of Management, McGill University, Montreal, QC H3A1G5, Canada. Goyenko may be reached at ruslan.goyenko@mcgill.ca and Sarkissian may be reached at sergei.sarkissian@mcgill.ca. We thank Yakov Amihud, Gurdip Bakshi, Kristin Forbes, Evan Gatev, David Goldreich, Denis Gromb, Andrew Karolyi, Michael Schill, Pedro Santa Clara, Rudi Schadt, Akiko Watanabe, and Guofu Zhou for useful comments. The paper has also benefited from the feedback of participants at the Seventh Darden International Finance Conference, the 2008 European Finance Association Meeting, the 2008 World Bank Conference on Risk Analysis and Management, the Bank of Canada Conference on Financial Market Stability, the 2009 Northern Finance Association Meeting, the 2009 Conference of the Society of Quantitative Analysts, as well as workshops at the ISCTE Business School, Durham University, and Luxembourg School of Finance. We are also grateful to Sam Henkel for providing international stock liquidity data. This paper was circulated earlier under "Flight to Liquidity and Global Equity Returns." Goyenko acknowledges financial support from IFM2. Sarkissian acknowledges financial support from IFM2 and SSHRC.

# Treasury Bond Illiquidity and Global Equity Returns


**ABSTRACT**

In this study, using data from 46 markets and a 30-year time period, we examine the impact of the illiquidity of U.S. Treasuries on global asset valuation. We find that it predicts stock market illiquidity as well as equity returns in both developed and emerging markets. Asset pricing tests further reveal that bond illiquidity is a priced factor even in the presence of other conventional risks. Since the illiquidity of Treasuries is known to reflect monetary shocks, our results suggest that it can be considered as one of the financial propagation channels of the U.S. monetary policy to other countries.



*JEL Classification*: E44; F36; G12; G15
*Keywords*: Bond illiquidity beta; Flight to liquidity; Illiquidity spillover; International asset pricing; Macroeconomic environment

## 1. Introduction

There is a well documented relation between monetary policy on the one side and stock and bond markets on the other. Fama and French (1989) find that dividend yield, default and term spreads are significant predictors of U.S. stock and bond returns. There is also substantial evidence on the impact of U.S. Treasury rates on expected returns in the U.S. and global equity markets.[1] Jensen, Mercer, and Johnson (1996) show that the main driving force behind the predictive power of these variables is the change in the monetary policy environment proxied by the Fed funds rate. Furthermore, such studies as Patelis (1997), Thorbecke (1997), Rigobon and Sack (2004), and Bernanke and Kuttner (2005) suggest not only predictive but also contemporaneous effect of U.S. monetary policy on stock returns.[2] They explain this finding by the persistent impact of U.S. monetary policy shifts on firm cash flows that last over several periods. Yet, Welch and Goyal (2008) cast doubts on the ability of interest rate and its various derivatives (e.g., term spread, default spread, etc.) to predict stock returns. In addition, a link between U.S. macroeconomic variables and foreign equity prices is not well established.[3]

In this paper, instead of analyzing the relation between Treasury rates or related U.S. interest-rate-based variables and stock markets around the world, we examine the relation between the *illiquidity* of Treasury bonds and international equity returns. Our choice of Treasury bond illiquidity is motivated by the following reasons.

First, there is substantial evidence on the importance of stock market illiquidity for equity returns in the U.S. (e.g., see Brennan and Subrahmanyam 1996; Pastor and Stambaugh, 2003; Acharya and Pedersen, 2005) as well as around the world (see Bekaert, Harvey, and Lundblad, 2007; Lee, 2011). While the effect of stock illiquidity on global equity returns is generally

[1] Fama, and Schwert (1977), Breen, Glosten, and Jagannathan (1989), Ang and Bekaert (2007), and Campbell and Thompson (2008) find strong predictive power of the U.S. T-bill rate for U.S stock returns. Campbell (1987) shows predictive power of the difference in returns of T-bills with different maturities. Harvey (1991), Ferson and Harvey (1993) and many others use U.S. T-bills as predictors of returns in the U.S. and world equity markets.
[2] Empirical support for the contemporaneous link between interest rates to stock returns is also presented in Stone (1974), Sweeny and Warga (1986), Ferson and Harvey (1993), and Scruggs (1998), among others.
[3] The few studies here are Bailey (1990) and Wongswan (2006, 2009) who document limited impact of the U.S. monetary policy proxy, the Federal Open Market Committee decisions, on equity markets in other countries.

accepted, we expect the effect of Treasury bond illiquidity on stock returns to be no less important. Indeed, Chordia, Sarkar, and Subrahmanyam (2005) and Baele, Bekaert, and Inghelbrecht (2010) document certain similarities between stock and bond market illiquidity. Furthermore, Goyenko and Ukhov (2009) observe that a distinctive feature of the illiquidity of Treasuries compared to that of stocks is that it reflects and transmits monetary policy shocks to equity markets. In addition, recent studies show that the change in the Fed fund rate is one of the main determinants of Treasury bonds' illiquidity (see Goyenko, Subrahmanyam, and Ukhov, 2009). Thus, given prior evidence of the Fed fund rate effect on stock and bond returns in the U.S. and U.S. stock illiquidity effect on global equity returns, we also expect the illiquidity of Treasuries to impact stock market returns around the world.

Second, U.S. Treasuries are typically viewed as the safest and most liquid asset class which comprises significant portion of diversified foreign equity portfolios. Investors outside the U.S. hold large and increasing stakes in U.S Treasuries. While at the end of 1996 they held close to 28% of all marketable Treasury securities outstanding, by the end of 2006 their holdings reached almost 45%.[4] This suggests that both foreign and domestic investors can be moving their funds in and out of Treasuries and thus affect Treasury market illiquidity (also see Longstaff, 2004; and Chordia, Sarkar, and Subrahmanyam, 2005). Therefore, similar to stock market illiquidity, Treasury bond illiquidity should also affect global stock returns. However, while the illiquidity effect related to stock trading costs should generally be subsumed by stock illiquidity (e.g., see Lee, 2011), the macroeconomic news component of Treasury illiquidity shocks should have an independent impact on equity prices.

We examine the impact of illiquidity of U.S. Treasuries on global equity returns using market-level data from 46 countries over the 30 year period from 1977 to 2006. This wide cross-sectional and time-series sample provides an ideal ground for analyzing the connection between

[4] Among these investors are of course financial institutions, such as central banks, that may have other needs rather than portfolio rebalancing. However, the holdings of foreign financial institutions in the T-bills, the asset of our primary interest, are decreasing over time suggesting that at least half of foreign holding in these securities belongs to non-financial institutional investors. Source: The Federal Reserve System, Treasury Bulletin, see http://www.ustreas.gov/tic/.

changes in the illiquidity of Treasuries and expected equity returns. If there is an illiquidity premium in asset returns associated with U.S. Treasuries, focusing on equities of both developed and emerging markets should result in particularly powerful tests and valuable cross-market evidence. Our main contribution is the finding of an economically and statistically significant illiquidity premium of U.S. Treasuries in global equity markets.

We proceed as follows. First, we find that Treasury bond illiquidity predicts global stock market illiquidity. In particular, an increase in U.S. Treasury bond illiquidity predicts higher stock market illiquidity both at the world and the local levels, even after controlling for various determinants of stock and bond market illiquidity, including the Fed funds rate. However, the reverse relation does not hold. These results suggest that U.S. monetary policy shocks are reflected initially in the illiquidity of Treasuries and then transferred into the illiquidity of global equity markets. Therefore, bond market illiquidity may be viewed a transmission channel of U.S. monetary policy shocks abroad.

Second, the literature on monetary policy effects on stock returns documents negative predictive and contemporaneous effects of monetary policy tightening on changes in share prices in the U.S. (e.g., see Jensen, Mercer, and Johnson, 1996; Patelis, 1997; Thorbecke, 1997; Bernanke and Kuttner, 2005). If the Treasury bond illiquidity is the transmission channel for U.S. monetary policy shocks to equity markets around the world, then we also expect it to have negative predictive and contemporaneous effects in international equity returns. We indeed find that bond illiquidity significantly negatively predicts stock returns, both in developed and emerging markets and in different sub-periods. This result is robust to the inclusion of other standard predictors of countries' equity returns such as local market returns, local dividend yields, the U.S. term spread, as well as local and world stock market illiquidity measures.

Finally, we explore the significance and magnitude of Treasury bond illiquidity risk in the setting of global asset pricing models. We first test a benchmark specification – a full-integration international market asset pricing model with two global risk factors: the world market portfolio return and the U.S. Treasury bond illiquidity factor. We then consider global pricing models that include the foreign exchange rate, the local equity market's variance and illiquidity, and the U.S. term spread as additional risk factors. Similar to Bekaert, Harvey, and Lundblad (2007), we conduct our estimation in two steps. In the first step, we use the multivariate GARCH (1,1) methodology and, for each country, compute the conditional return variance and the set of conditional covariances between local stock market returns and the model-specific risk factors. In the second step, we use GMM and estimate prices of risk for both the entire sample of countries and for developed and emerging market subsamples. Since the contemporaneous covariance between bond illiquidity and stock returns is also negative, our asset pricing tests, as expected, show a negative and significant price of U.S. Treasury bond illiquidity risk. This implies that bond illiquidity risk is associated with positive premium in global equity markets.[5] This result holds in the presence of other world and local risk factors considered, for the full sample of countries as well as for the developed and emerging market subsamples.

The estimates of the price of bond illiquidity risk are usually larger in magnitude in emerging markets and in such countries as Greece and Portugal, which were classified as developed in the later part of our sample. This is natural as those markets are more exposed to term/default spread or interest rate shocks than well developed countries. In our benchmark model, in economic terms, the average annual premium for bond illiquidity risk is between 1.0% and 1.3%. This is comparable in magnitude to the stock illiquidity premium of 1.1% per annum reported by Acharya and Pedersen (2005) for the U.S. equity market. The only other consistently priced factor across all models, not surprisingly, is the world market portfolio return.

Our study uses market-level rather than individual security returns for a variety of reasons. During economic uncertainty or market downtrends, which usually witness an increased desire among investors for U.S. Treasuries, correlations across risky assets increase.[6] Yet, the continuing benefits of international diversification, as shown in Ang and Bekaert (2002) and

[5] Positive illiquidity premium is consistent with the observation of Patelis (1997) that even though the effect of monetary policy is negative and persistent, it decreases over time suggesting that expected returns are increasing after monetary policy contraction shock.
[6] See Longin and Solnik, (1995, 2001), De Santis, and Gerard (1997), and Ang and Chen (2002).

many other papers, imply that this increase in return correlations is larger within a country than across countries (e.g., recall the Asian crisis of 1997, the Russian crisis of 1998, or the Greek crisis of 2010). Besides, dealing with firm returns in a conditional setting such as ours is not feasible from a computational viewpoint. Finally, there are also global data limitation issues: for instance, we cannot obtain reliable measures of stock market illiquidity at the daily or weakly frequencies.

The main contribution of our study is showing that U.S. monetary policy shocks are transmitted to equity prices around the world. This finding is related to the literature on "financial propagation" mechanisms as well as the "credit channel" of monetary policy transmission (e.g., see Bernanke and Blinder, 1988, 1992; Romer and Romer, 1990; Bernanke and Gertler, 1995; Kashyap and Stein, 2000). Our results suggest that the illiquidity of Treasuries can be considered as one of the strong financial propagation channels of shifts in the U.S. monetary policy to other countries.

The rest of the paper is organized as follows. Section 2 describes the data. Section 3 offers the initial analysis on the importance of the U.S. Treasury bond illiquidity for global equity markets. In particular, we analyze the relation between bond illiquidity and global stock market illiquidity (at the world and local levels) as well as examine predictive regressions of stock market returns on lagged values of bond illiquidity and other variables. In Section 4, we develop our conditional asset pricing methodology. Section 5 presents the results of asset pricing tests. In this section, we also relate our estimates of the bond illiquidity risk to a set of country-level macroeconomic and financial variables. In Section 6, we offer robustness tests. Section 7 concludes.

## 2. Data

Our data sample consists of 46 countries, of which 23 are classified as developed and 23 as emerging. The sample covers the 30-year period from January 1977 to December 2006, although for many countries the time-series data start significantly later than 1977. For each country, we collect monthly local equity market returns in U.S. dollars and dividend yields from Datastream (for developed markets) and IFC Global Indices (for emerging markets). We construct excess returns by subtracting the one-month U.S. Treasury bill rate from gross returns. Following Bekaert, Harvey, and Lundblad (2007) and Lee (2011), our proxy for stock market illiquidity in each country is the zero-return measure (Zeros) suggested by Lesmond, Ogden, and Trzcinka (1999). This measure is motivated by data limitations, which are especially pronounced in emerging markets. Note, however, that Zeros is directly related to trading volume. More illiquid stocks have less frequent trading and therefore a higher incidence of zero returns.[7] We use the value-weighted proportion of zero daily returns across all firms in a country during a month. World stock market illiquidity is the value-weighted average of country-level aggregate illiquidity series.

Goyenko, Subrahmanyam, and Ukhov (2009) analyze the illiquidity of U.S. Treasuries across all maturities and on-the-run/off-the-run status and find that the illiquidity of off-the-run T-bills with maturities of up to one year best captures the illiquidity of the Treasury market overall. Accordingly, we use the illiquidity of off-the-run T-bills as our proxy for the illiquidity of the U.S. Treasury bond market. More specifically, we use the average percentage bid-ask spread of off-the-run U.S. T-bills with maturities of up to one year to proxy for U.S. Treasury bond market illiquidity. The quoted bid and ask prices come from CRSP's daily Treasury Quotes file. This file includes Treasury fixed income securities of three and six months, as well as 1, 2, 3, 5, 7, 10, 20, and 30 years, to maturity. Under the standard definition, when a new security is issued it is considered to be on-the-run and the older issues are treated as off-the-run. We use the quotes for three-, six-, and 12-month securities. For each month the monthly average spread is first computed for each security as the average proportional daily spread for the month and then

[7] Fong, Holden, and Trzcinka (2009) find that Zeros efficiently captures the time-series patterns of stock market liquidity compared to effective spread-based benchmarks. They analyze monthly data across 39 countries over the 1996-2007 period.

equally weighted across short-term assets.[8] These data have also been used by Acharya, Amihud, and Bharath (2009) and Baele, Bekaert, and Inghelbrecht (2010). The primary motivation for using the CRSP data is to have a long enough Treasury bond illiquidity time series to be able to study the connection between economic environment, liquidity conditions, and equity prices across different countries; to our knowledge, CRSP is the only data source that allows for the use of a sufficiently long period to subsume a variety of economic events. However, in robustness tests we use a bond illiquidity measure estimated from high-frequency intraday GovPX data that starts in the 1990s.

Table 1 shows the number of observations, means, volatilities, and first-order autocorrelations of monthly excess equity returns, dividend yields, and the stock liquidity measure for each country and for the world market. The number of observations corresponds to the equity market returns. Not surprisingly, the average monthly returns and volatilities in emerging markets are higher than those in developed markets. The autocorrelation of dividend yields is very high, in excess of 0.90 in all but five countries. Market illiquidity based on the zero-return measure is also higher on average in emerging markets than developed, as expected. Zeros is highly correlated with transaction costs, but it does not directly indicate the magnitude of illiquidity (see Hasbrouck, 2009; Goyenko, Holden, and Trzcinka, 2009). Rather, this measure gives us only a relative sense of the magnitude of illiquidity. For example, the U.S., which is the most liquid market, has the lowest realization of Zeros at 8.6%, whereas emerging markets, which are perceived to be more illiquid, observe Zeros in the range of 20% and 50%.

Surprisingly, the average illiquidity in the U.K. is 49.3%. It is similar in magnitude to the estimate reported by Lee (2011). In his sample, which starts in 1988, the average zero-return proportion is 37.8%. The higher magnitude of our estimate can be related to the earlier sample period that we use, from 1977. Lee (2011) suggests that higher proportion of zeros is a common

[8] Our results are similar when non-scaled (raw) quoted spreads are used as an alternative to proportional quoted spreads. This is consistent with Chordia, Sarkar, and Subrahmanyam (2005), who show that the daily correlation between quoted and effective spread changes in the bond market is 0.68 over their nine-year sample period. Thus, quoted spreads are reasonable liquidity proxies.

characteristic for the U.K. market across all size portfolios and is not driven by outliers. The zero-return measure also shows autocorrelation but not to the same extent as dividend yields. The only country with a negative first-order autocorrelation of illiquidity is China.

## 3. Preliminary Analysis

### *3.1. Treasury Bond and Stock Market Illiquidity*

We first investigate the relation between U.S. Treasury bond illiquidity and stock market illiquidity around the world. This is important since if illiquidity of one market (e.g., Treasuries) affects illiquidity of the other market (e.g., equity), it should also forecast the other market's illiquidity premium. It is known that shocks to bond illiquidity associated with the U.S. monetary policy shifts impact U.S. stock market illiquidity (Goyenko and Ukhov, 2009). Therefore, we want to see whether there exists a similar propagation mechanism between the U.S. Federal Reserve actions and the illiquidity of international equity markets. The results are reported in Table 2.

In Panel A of Table 2, the world stock market illiquidity, $L_{w,t}$, is regressed on the lagged Treasury bond illiquidity, $L_{B,t-1}$, with and without control variables. We report point estimates and robust t-statistics based on the Newey-West correction for six lags of the standard error. The first two lags of $L_{w,t}$ are included in each regression to control for persistence in the series, but their coefficients are not reported. Regression (1) shows the base specification between $L_{B,t-1}$ and, $L_{w,t}$. We find that the coefficient on bond illiquidity is positive and significant at the 5% level, indicating that an increase in bond illiquidity leads to an increase in world stock market illiquidity next period. This result suggests the presence of illiquidity shock spillover effects between the two markets.

Prior studies suggest that returns and volatility of returns are important drivers of illiquidity (see Amihud and Mendelson, 1986; Benston and Hagerman, 1974). Therefore, in Regression (2) of Panel A, we add to the above specification controls for the world stock market's lagged return, $r_{w,t-1}$, and volatility, $\sigma_{w,t-1}$.[9] The results show that lagged bond illiquidity remains statistically significant. In contrast, the control variables, which have been shown by Chordia, Roll, and Subrahmanyam (2001) and Chordia, Sarkar, and Subrahmanyam (2005) to predict stock market illiquidity in the U.S., do not appear to affect world stock market illiquidity.

As mentioned earlier, the Treasury bond illiquidity captures a substantial portion of monetary policy shocks (see Goyenko, Subrahmanyam, and Ukhov, 2009). In turn, monetary policy can directly affect stock market illiquidity by tightening the inventory constraints of market makers and increasing the borrowing costs of trading (see, e.g., Chordia, Sarkar, and Subrahmanyam, 2005; Hameed, Kang, Viswanathan, 2010). Furthermore, money market fund flows and consumer confidence can impact bond market illiquidity premiums (Longstaff, 2004). Due to the linkages between the bond and stock markets, these variables may also affect stock market illiquidity.

Regressions (3) and (4) of Panel A therefore consider two monetary policy controls (the lagged change in the Fed funds rate, $FED_{t-1}$, and the lagged term spread, $TERM_{t-1}$) and two controls based on Longstaff (2004) (the lagged percentage change in the amount of funds held in money market mutual funds, $MMF_{t-1}$, and the lagged change in the consumer confidence index, $CCI_{t-1}$), respectively.[10] The results show that current-period bond illiquidity continues to predict next-period world stock market illiquidity at the 10% level or better. Moreover, consistent with Amihud and Mendelson (1986) and Benston and Hagerman (1974), world market volatility has positive and marginally significant power to predict world stock market illiquidity. Note that the slope on $FED_{t-1}$ is only marginally significant in Regression (3), which suggests that bond illiquidity has more important direct effect on stock market illiquidity than monetary policy.

Regression (5) of Panel A includes all control variables above except MMF due to its high correlation with the term spread. The results again show a positive and significant link between

[9] The monthly stock market volatility for each market in a given month is computed as the standard deviation of daily returns in that market and month. Daily returns are again from Datastream and IFC.

[10] The term spread is the difference in yields between the 10-year U.S. Treasury note and the one-month T-bill. Data on the amount of funds held in money market mutual funds come from the Federal Reserve Board, and data on the consumer confidence index, which is divided by 100, come from the Conference Board.

lagged bond illiquidity and world stock market illiquidity, and lagged world market volatility is again only marginally significant.

Next, Panel B of Table 2 shows the results of regressing individual sample countries' stock market illiquidity on Treasury bond illiquidity. To properly address cross-country correlations and substantial persistence in stock market illiquidity series (see Table 1), we use a structural dynamic panel data estimation technique based on Arellano and Bover (1995) and Blundell and Bond (1998, 2000). This technique is a GMM procedure that allows one to estimate panel data taking into account serial and cross-sectional correlation (via time effects), heteroskedasticity, as well as endogeneity of some explanatory variables. The estimator is based on a system of moment conditions that contain not only original equations but also first-differenced equations.[11] This panel regression model can be written as follows:

$$L_{i,t} = \alpha_0 + \sum_{l=1}^{2} \phi_l L_{i,t-l} + \beta X_{it-1} + f_i + d_t + e_{i,t} \, , \qquad (1)$$

where the vector of independent variables, *X*, is either a scalar, $L_B$, or $L_B$ augmented by a subset of control variables {$r_i$, $\sigma_i$, FED, TERM, MMF, CCI}, depending on the regression specification, $f_i$ captures country-specific effects, and $d_t$ captures calendar effects. The first two lags of $L_{i,t}$ are included in each estimation, but their coefficients are not reported.

In Regression (1) of Panel B we observe that, in line with our results in Panel A, bond illiquidity significantly positively predicts stock illiquidity at the country (i.e., local) level. This relation remains intact after adding various control variables in Regressions (2)-(5). Consistent with the inventory paradigm (see, e.g., Ho and Stoll, 1983; and O'Hara and Oldfield, 1986), we now find that local stock market volatility positively and significantly (at the 10% level or better) predicts local stock market illiquidity in all regression specifications. Changes in the Fed funds rate also positively affect local stock market illiquidity, similar to our findings in Panel A. We

[11] Note that Arellano and Bover's (1995) procedure is commonly applied to panel data when it is not possible to run vector autoregression analysis (VAR) due to cross-sectional correlations. Essentially, this procedure gives robust VAR estimates for unbalanced panel data.

also find a strong effect of lagged changes in the U.S. consumer confidence index, CCI, on stock market volatility across our sample countries.

Finally, Panel C of Table 2 shows the results of the reverse relation, that is, of the predictive effect of world stock market illiquidity for Treasury bond illiquidity. An increase in stock market illiquidity may result in increased flows of funds into Treasuries (flight-to-liquidity), reducing the illiquidity of Treasury bonds. Stock illiquidity may thus have a negative impact on next-period bond illiquidity. Other variables may also have predictive power for Treasury bond illiquidity. In Panel C, we include the same control variables as those used in Panel A. Also as in Panel A, the first two lags of $L_{B,t}$ are included in each regression (their coefficients are not reported), and we again use t-statistics based on the Newey-West correction for six lags of the standard error. We find that across all specifications world stock market illiquidity has no significant predictive effect on Treasury bond market illiquidity. Volatility of the world stock market, however, has a negative and marginally significant effect on bond illiquidity, that is, an increase in stock market volatility increases flows of funds into U.S. Treasuries, improving their liquidity. This is consistent with flight-to-liquidity episodes. We also find, similar to Goyenko, Subrahmanyam, and Ukhov (2009), that changes in the Fed funds rate have positive predictive power for bond illiquidity. Taken together, the results suggest that both U.S. monetary policy and world stock market volatility impact U.S. Treasury market illiquidity. The illiquidity of Treasuries increases in response to monetary policy tightening and decreases in response to fund outflows from global stock markets.

In sum, Table 2 shows that Treasury bond illiquidity has predictive power for stock market illiquidity. In particular, an increase in bond illiquidity predicts an increase in both world and country-specific stock market illiquidity. The reverse relation, however, does not hold. These findings reveal that illiquidity shocks related to changes in the U.S. monetary policy are reflected first in the illiquidity of the U.S. Treasury market, an important source of immediate liquidity provision, and then transferred into the illiquidity of stock markets around the world.

### *3.2. Predictive Regressions of Equity Returns*

Given the evidence that Treasury bond illiquidity predicts global stock market illiquidity, in this section, we test whether it has predictive power for global equity returns as well. Since a positive shock to bond illiquidity is associated with tightening of the U.S. monetary policy, and the effect of the latter on expected stock returns is negative (see Patelis, 1997; Thorbecke, 1997; and Bernanke and Kuttner, 2005), we also expect a *negative* relation between bond illiquidity and expected equity returns.

Note that both bond illiquidity and stock market illiquidity are persistent. Ferson, Sarkissian, and Simin (2003) warn against using standard statistical inference in regressions of stock returns on lagged instruments when the regressors are autocorrelated. Therefore, to preclude concerns about spurious regression biases, in the subsequent analysis we follow Pastor and Stambaugh (2003) and Acharya and Pedersen (2005) and use the AR(2) residuals as an illiquidity measure of both the Treasury bond and global stock markets. To reduce the impact of outliers on our estimation results, we winsorize bond and stock market illiquidity shocks at the 1st and 99th percentiles. Table 3 presents test the results of predictive regressions for global and local excess market returns. The two control variables included in all panels are the U.S. term spread and the January dummy; the latter variable is included in every regression.

Panel A of Table 3 reports the results for the world equity market return. In particular, the panel reports point estimates and robust t-statistics based on the Newey-West correction for six lags of the standard error. The regressions include as global stock market controls lagged values of the world market's return, illiquidity, and dividend yield. We conduct our estimation on both the full sample period (columns 1-3) and the three ten-year subperiods of 1977-1986, 1987-1996, and 1997-2006 (columns 4-6). The first three columns show that the slope on bond illiquidity is negative and significant at the 1% level, consistent with our expectations. None of the other variables is significant. The last three columns show that the negative relation observed between lagged bond illiquidity and stock returns is present in each of the subperiods, with its magnitude increasing towards later years of the sample.

The predictive relation between Treasury bond illiquidity and world equity market excess returns is economically important as well. Since one standard deviation of bond illiquidity is 0.002, a one-standard deviation positive shock to bond illiquidity, based, for instance, on Regression (3) output implies a decrease in next-period world market excess returns of -3.091 times 0.002. This amounts to a return decline of 62 basis points per month.

Panel B of Table 3 reports panel regression results for local stock market returns. Our controls now include country-level lagged values of equity market returns, illiquidity, and dividend yields. To account for cross-market correlations and average country-specific characteristics, all regressions include country and year fixed effects, and we cluster standard errors by month. Again, columns 1-3 correspond to full sample period tests, while columns 4-6 correspond to the sub-period tests. The first three regressions show that over the entire sample period, bond illiquidity retains its negative and statistically significant predictive power for local stock returns. Moreover, this relation mostly survives the subperiod tests.[12] Across all regression specifications, the coefficients on $L_B$ are comparable in magnitude to those in Panel A. Another variable which often shows significance in predictive tests is the local dividend yield. There is no systematic evidence on the importance of lagged stock market illiquidity for equity returns.

In Panel C of Table 3, we split the sample countries into 23 developed and 23 emerging markets and repeat the first three tests of Panel B. Columns 1-3 report the estimation results for the developed markets, while columns 4-6 report the results for the emerging markets. The slope on lagged bond illiquidity is negative and significant at the 5% level or better across all six specifications. However, its magnitude for emerging markets is more than four times larger than that for developed markets. Thus, emerging markets, which tend to be less liquid, experience stronger illiquidity effects. This is consistent with the U.S. evidence that monetary policy effects are stronger for smaller, more illiquid stocks. Dividend yields predict stock returns in developed

[12] In the last sub-period (1997-2006, column 6) the effect of bond illiquidity is insignificant at conventional significance levels. However, over the 1998-2006 period, which excludes the highly volatile returns of emerging markets during the Asian crisis, the t-statistic associated with $L_{B,t-1}$ becomes significant again, reaching the value of -2.30.

markets only. Interestingly, lagged local market illiquidity is negative and significant for developed countries, but essentially zero for emerging markets.[13]

Overall, Treasury bond illiquidity predicts global stock returns at both the world and the individual country levels, over different sub-periods, and across developed versus emerging markets. This result, which is statistically and economically significant even after controlling for common predictors of equity returns and stock market illiquidity, points out that changes in the U.S. monetary policy affect not only stock prices in the U.S. but also overseas equities. In the next section, we investigate the main pricing implications of bond illiquidity for global equity returns.

## 4. Conditional Methodology

### *4.1. General Framework*

In this section, we test four asset pricing models of global equity returns under full and partial market integration. All models use Treasury bond illiquidity as a proxy for the propagation channel of U.S. monetary policy shifts.[14] We assume constant prices of all risk factors.

Model I. If country *i* is integrated with the world and purchasing power parity holds across countries, then country *i*'s expected return at time *t* given the information available at time *t*-1 is determined by its conditional covariances with the return on the world market portfolio and with Treasury bond illiquidity, that is,

$$E_{t-1}(r_{i,t}) = \lambda_w \,\mathrm{Cov}_{t-1}(r_{i,t}, r_{w,t}) + \lambda_{LB} \,\mathrm{Cov}_{t-1}(r_{i,t}, L_{B,t}), \qquad (2)$$

where $\lambda_w$ is the price of world market risk and $\lambda_{LB}$ is the price of the Treasury bond illiquidity risk. Equation (2) is our benchmark two-factor "World CAPM" that we call WCAPM-LB.

[13] Bekaert, Harvey, and Lundblad (2007) find a significantly positive (negative) relation between excess returns in closed (open) emerging markets and lagged local stock market illiquidity. This implies a generally flat relation between lagged stock liquidity and excess returns in emerging markets over the full sample period, similar to our result. Also note that their relation between stock market liquidity and excess returns in open emerging markets resembles ours in developed markets, as one would expect from liberalized economies.
[14] Note that since the Treasury bond illiquidity risk is a global factor it cannot be present in fully segmented markets.

Economically and statistically significant $\lambda_{LB}$ would suggest that the risk associated with changes in the U.S. monetary policy is priced in global markets. Strictly speaking though, significant $\lambda_{LB}$ will be associated with U.S. monetary shifts only in the presence of stock illiquidity risk in the asset pricing model (see Model IV below). In the absence of stock illiquidity, due so certain commonality between bond illiquidity and stock illiquidity, bond illiquidity may also capture risks embedded in the trading costs of equities.

Note that contemporaneous effect of monetary policy tightening on equity returns is generally negative (see Thorbecke, 1997; and Bernanke and Kuttner, 2005). Therefore, similar to a negative predictive relation, we also expect a *negative* contemporaneous relation between bond illiquidity and global stock returns. This effect will also be similar to that between stock illiquidity and equity returns (see Amihud, 2002), and imply negative on average $Cov_{t-1}(r_{i,t}, L_{B,t})$ term. Therefore, if bond illiquidity is a systematic risk factor in international equity markets, $\lambda_{LB}$ must have negative sign as well. This is our main testable hypothesis.

We must point out that empirical literature documents that another financial variable closely related to monetary policy, the short-term interest rate, also has negative predictive and contemporaneous effects on stock prices (e.g., see Breen, Glosten, and Jagannathan, 1989; Fama and Schwert, 1977; Campbell, 1987). However, Bernanke and Kuttner (2005) point out that the reaction of equity prices to monetary policy is not directly related to the policy's impact on the real interest rate.

Model II. If there are deviations in purchasing power parity across countries, then exchange rate risk may also be priced (see Dumas and Solnik, 1995). Model II extends Model I to accommodate this factor as follows:

$$E_{t-1}(r_{i,t}) = \lambda_w \operatorname{Cov}_{t-1}(r_{i,t}, r_{w,t}) + \lambda_{LB} \operatorname{Cov}_{t-1}(r_{i,t}, L_{B,t}) + \lambda_c \operatorname{Cov}_{t-1}(r_{i,t}, r_{c,t}), \tag{3}$$

where $r_{c,t}$ is the return on the currency basket deposit at time $t$ and $\lambda_c$ is the price of currency risk. In our estimations, the return on the currency basket deposit is calculated as the equally weighted average change in exchange rates between the U.S. dollar and four global currencies: the British Pound, Euro, Japanese Yen, and Swiss Franc.[15]

Model III. A country may not be fully integrated with the world. Errunza and Losq (1985) develop a model where expected return on a risky security in such country is determined by a global risk premium and an additional risk premium proportional to the country's conditional market risk. If country *i* is fully segmented, its expected return at time *t* given the information available at time *t*-1 is based only on its conditional variance with the market returns, i.e., $E_{t-1}(r_{i,t}) = \lambda_i \mathrm{Var}_{t-1}(r_{i,t})$, where $\lambda_i$ is the price of country *i* risk. We combine this term with Model I, following similar econometric specifications of Chan, Karolyi, and Stulz (1992), Bekaert and Harvey (1995), De Santis and Gerard (1997), and many others and obtain an asset pricing model of partial world market integration.

$$E_{t-1}(r_{i,t}) = \lambda_w \mathrm{Cov}_{t-1}(r_{i,t}, r_{w,t}) + \lambda_{LB} \mathrm{Cov}_{t-1}(r_{i,t}, L_{B,t}) + \lambda_i \mathrm{Var}_{t-1}(r_{i,t}). \quad (4)$$

In this model, the expected return in country *i* is determined based on its conditional covariances with the two global risk factors as well as respective country risk.

Model IV. Recent research shows that stock market illiquidity is an important factor for U.S. stock returns (see, e.g., Amihud, 2002; Pastor and Stambaugh, 2003; Acharya and Pedersen, 2005). There is some evidence that stock market illiquidity is also important in global markets (e.g., Bekaert, Harvey, and Lundblad, 2007; Lee, 2011). To control for stock market illiquidity, we further extend the partial integration model (Model III) to include the second country-specific factor. This yields the following model

$$E_{t-1}(r_{i,t}) = \lambda_w Cov_{t-1}(r_{i,t}, r_{w,t}) + \lambda_{LB} Cov_{t-1}(r_{i,t}, L_{B,t}) + \lambda_i Var_{t-1}(r_{i,t}) + \lambda_{Li} Cov_{t-1}(r_{i,t}, L_{i,t}), \quad (5)$$

where $\lambda_{Li}$ is the price of equity market illiquidity risk in country *i*.

[15] Across various currency pairs, we observed that only changes in the JPY/USD rate were significantly related to world stock market excess returns. Since the trade- or GDP-weighted approach assigns the JPY/USD rate a weight of less than 15% over the sample period, we account for the JPY/USD rate to the larger degree by using the equally weighted setting (25% for each currency pair). Note, however, that replacing our currency basket with individual exchange rates does not materially impact our test results.

Note that it is possible to combine Models II and IV, which would result in a five-factor model, but we do not pursue this setting due to the added estimation difficulties. Following Acharya and Pedersen (2005), one could also consider other stock market illiquidity based covariance risks, such as $Cov_{t-1}(r_{w,t}, L_{i,t})$, $Cov_{t-1}(L_{B,t}, L_{i,t})$, or $Cov_{t-1}(L_{w,t}, L_{i,t})$. However, these additions will again render our estimation impractical.

### *4.2. Estimation Details*

Evaluating Models I through IV jointly across 46 countries in a conditional framework with unknown conditional variances and covariances is practically impossible. We therefore estimate our asset pricing models in two steps. While the two-step estimation framework is usually associated with an errors-in-variables problem, it is often the only technique for testing multi-country or multi-asset conditional asset pricing models.[16]

In the first step, we estimate the conditional variances of equity market returns and their covariances with all risk factors depending on the model specification. We obtain these estimates separately for each country within a multivariate GARCH (1,1) setting that includes return and risk factor dynamics. We follow Harvey (1991), Ferson and Harvey (1993), and many others and model country equity returns and risk factors as linear functions of global and local information variables.

The choice of our information variables is determined by previous literature and the results in Tables 2 and 3. First, for the local (world) market return, we use the first lags of the local (world) market return, local (world) dividend yield, U.S. term spread, Treasury bond illiquidity, as well as local (world) stock market illiquidity. We include the lagged bond illiquidity and stock market illiquidity based on our Table 2 and other studies (e.g., Bekaert, Harvey, and Lundblad, 2007), respectively. Including lagged stock market returns is a common

[16] For example, Bekaert, Harvey, and Lundblad (2007) model stock market liquidity in emerging countries using a two-step estimation procedure, where the first step is based on the VAR(1) framework and the second on GMM. Engle (2002) examines conditional correlations across multiple assets using a two-step approach with multivariate GARCH models.

practice in conditional asset pricing, although they are often insignificant.[17] Second, for bond illiquidity, the instruments are lagged stock market volatility and the change in the Fed funds rate, which come from our Table 2 and Goyenko, Subrahmanyam, and Ukhov (2009). Third, the change in the exchange rate is predicted by the lagged world market return and the one-month Eurodollar deposit rate, following Dumas and Solnik (1995). In unreported results we find that the Eurodollar rate predicts changes in our worldwide exchange rate.[18] Finally, stock market illiquidity is predicted by lagged values of bond illiquidity, stock market return, and volatility. This choice is based on our results in Table 2 as well as extant studies (see Chordia, Roll, and Subrahmanyam, 2001; Chordia, Sarkar, and Subrahmanyam, 2005).

Based upon the discussion above, for our Model I (WCAPM-LB) and Model III we initially estimate the following trivariate GARCH (1,1) system for each country:

$$r_{i,t} = \delta_{10} + \delta_{11} L_{B,t-1} + \delta_{12} r_{i,t-1} + \delta_{13} L_{i,t-1} + \delta_{14} DY_{i,t-1} + \delta_{15} TERM_{t-1} + e_{i,t} \tag{6a}$$

$$r_{w,t} = \delta_{20} + \delta_{21} L_{B,t-1} + \delta_{22} r_{w,t-1} + \delta_{23} L_{w,t-1} + \delta_{24} DY_{w,t-1} + \delta_{25} TERM_{t-1} + e_{w,t} \tag{6b}$$

$$L_{B,t} = \delta_{30} + \delta_{31} \sigma_{w,t-1} + \delta_{32} FED_{t-1} + e_{LB,t} . \tag{6c}$$

For Model II, we add the relation that governs the dynamics of currency returns,

$$r_{c,t} = \delta_{40} + \delta_{41} r_{w,t-1} + \delta_{42} Euro\$_{t-1} + e_{c,t} \ , \tag{6d}$$

while for Model IV we add instead the predictive relation for local stock market illiquidity,

$$L_{i,t} = \delta_{50} + \delta_{51} L_{B,t-1} + \delta_{52} r_{i,t-1} + \delta_{53} \sigma_{i,t-1} + e_{Li,t} . \tag{6e}$$

[17] We use the lagged dividend yield and term spread following Fama and French (1989), who observe that these variables predict stock returns. Our results in Table 2, while not showing significance for the term spread, show significant predictive power of dividend yields (at least using the standard statistical inference). We include lagged bond illiquidity and stock market illiquidity based on our Table 2 and other studies (e.g., Bekaert, Harvey, and Lundblad, 2007), respectively. Including lagged stock market returns is a common practice in conditional asset pricing, although they are often insignificant.

[18] Ideally, we would like to have short-term rates for all the currencies contributing to our currency basket. However, as Dumas and Solnik (1995) note, expanding the instrument set to include several interest rates can quickly worsen the finite-sample properties of estimates due to high auto- and cross-correlations of interest rates.

We also estimate system (6) for the world market portfolio. In this case, equation (6a) is dropped, and, for Model IV, all local market variables in equation (6e) are replaced with their corresponding world market characteristics, that is

$$L_{w,t} = \delta_{50} + \delta_{51} L_{B,t-1} + \delta_{52} r_{w,t-1} + \delta_{53} \sigma_{w,t-1} + e_{Lw,t} \, . \qquad (6f)$$

In the full system of equations (6a-f), the error term is $e_t = [e_{i,t}, e_{w,t}, e_{c,t}, e_{LB,t}, e_{Li,t}, e_{Lw,t}]$. It is assumed to be a multivariate normal distribution with conditional variance-covariance matrix $H_t$. The matrix $H_t$ has the BEKK structure ensuring that it is parsimonious and positive definite (see Engle and Kroner, 1995):

$$H_t = C'C + A'e_{t-1}e'_{t-1}A + B'H_{t-1}B \, ,$$

where $C$ is an (MxM) upper triangular matrix and A and B are (MxM) diagonal matrices, where $M$ is the number of equations being estimated under different model specifications. We therefore assume that current-period variance depends only on lagged conditional variance and lagged squared errors, while current-period covariance depends only on lagged covariance and the lagged cross-product of errors. Similar specifications are used in Bekaert and Harvey (1995), DeSantis and Gerard (1997), and other papers. To obtain the parameter estimates, we employ the Berndt, Hall, Hall, and Hausman (BHHH) optimization algorithm. We use conditional covariances between $r_{i,t}$ on the one side and $r_{w,t}$, $L_{B,t}$, $r_{c,t}$, and $L_{i,t}$ on the other, as well as conditional variances of $r_{i,t}$ for each country obtained from the system (6) for the second- step GMM estimation.

In the second step, we use panel GMM and estimate pricing moment conditions across all countries (or country groups) and the world market. For example, the moment conditions for Model IV are:

$$\begin{aligned} \zeta_{i,t} &= r_{i,t} - \lambda_w \hat{C}ov_{t-1}(r_{i,t}, r_{w,t}) - \lambda_{LB} \hat{C}ov_{t-1}(r_{i,t}, L_{B,t}) - \lambda_i \hat{V}ar_{t-1}(r_{i,t}) - \lambda_{Li} \hat{C}ov_{t-1}(r_{i,t}, L_{i,t}) \\ \zeta_{w,t} &= r_{w,t} - \lambda_w \hat{V}ar_{t-1}(r_{w,t}) - \lambda_{LB} \hat{C}ov_{t-1}(r_{w,t}, L_{B,t}) - \lambda_c \hat{C}ov_{t-1}(r_{w,t}, r_{c,t}) - \lambda_{Lw} \hat{C}ov_{t-1}(r_{w,t}, L_{w,t}) \end{aligned} , \qquad (7)$$

where $\zeta_{i,t}$ and $\zeta_{w,t}$ are the error terms of the country *i* and world market excess return equations at time *t*, respectively, *i=1,…N*, and *N* is the number of countries (46 for the whole sample or 23 for the sub-samples of developed and emerging markets). The "hat" indicator denotes the conditional variances and covariances from the multivariate GARCH (1,1) estimation. At this stage, we compute the following prices of risk:

Model I: $\lambda_w$, $\lambda_{LB}$;

Model II: $\lambda_w$, $\lambda_{LB}$, $\lambda_c$;

Model III: $\lambda_w$, $\lambda_{LB}$, $\lambda_i$, *i=1,…N*;

Model IV: $\lambda_w$, $\lambda_{LB}$, $\lambda_{Lw}$, $\lambda_i$, $\lambda_{Li}$, *i=1,…N*.

To create orthogonality conditions in an overidentified yet parsimonious system, we use instruments that can be implemented with various asset pricing models. This approach facilitates comparison of test results across models. Our most commonly used instrument vector $Z$, which is largely motivated by the predictive regression results in Table 3, includes a constant and three global information variables, namely, the lagged values of Treasury bond illiquidity, the world market portfolio return, and the world dividend yield, that is,

$$Z_{t-1} = [1, L_{B,t-1}, r_{w,t-1}, DY_{w,t-1}]. \tag{8}$$

This gives a total of (4N+4) orthogonality conditions in the GMM estimation. However, in smaller GMM systems (e.g., the WCAPM-LB specification, Model I), we also use a shorter instrument vector by dropping the world lagged dividend yield from (8), while in the larger systems (e.g., Models II-IV), we also use an alternative instrument set in which the lagged world dividend yield in (8) is replaced with the world stock market illiquidity shocks and the U.S. term spread. These variations allow us to examine the sensitivity of our results to the instrument choice.

Following the studies of GMM performance in small samples (Andersen and Sørensen, 1996; Ferson and Foerster, 1994), we use Bartlett kernel, Andrews' bandwidth, and iterative updating of both the weighting matrix and the coefficients in all our GMM estimations. Furthermore, to facilitate convergence, we apply the pre-whitening of the weighting matrix as suggested by Andrews and Monahan (1992).[19]

# 5. Empirical Tests

### *5.1. Conditional Treasury Bond Illiquidity Betas*

We start by examining the outcome of our multivariate GARCH (1,1) model based on equations (6a-c).[20] In particular, given the estimates of the conditional variance of Treasury bond illiquidity, $\hat{\mathrm{V}}\mathrm{ar}_{t-1}(L_{B,t})$, and the conditional covariance of country returns with bond illiquidity, $\hat{\mathrm{C}}\mathrm{ov}_{t-1}(r_{i,t}, L_{B,t})$, we can construct for each country *i* the conditional bond illiquidity beta as:

$$Beta_{i,t-1}(L_{B,t}) = \hat{C}ov_{t-1}(r_{i,t}, L_{B,t}) / \hat{V}ar_{t-1}(L_{B,t}). \tag{9}$$

In Figure 1, we plot the time series of the conditional Treasury bond illiquidity beta. The figure depicts the average betas for developed markets (Plot A) and emerging markets (Plot B). These betas are averaged for each month across 23 developed and 23 emerging markets, respectively. We can see that the betas for both country groups are highly volatile, especially after 1987. The average conditional bond illiquidity beta for developed markets is close to but less than zero, while that for emerging markets is much more negative (larger in absolute terms). This result is consistent with the intuition that the effect of U.S. monetary policy tightening should be more pronounced in emerging markets. Indeed, firms in these countries are subject to

---

[19] Andersen and Sørensen (1996) find that an estimation using a fixed number of lags in the weighting matrix is inferior to one using an automatic (data-dependent) bandwidth, such as Andrews' bandwidth. They also find that the standard Bartlett kernel estimator is superior to the quadratic one in many model specifications and that pre-whitening can often be helpful in the estimation when the sample size is relatively small. Ferson and Foerster (1994) show that iterated GMM has better finite sample properties than the standard two-stage approach.

[20] In unreported results, we compute the Jarque-Bera statistic (JB) for normality of residuals and compare it to that of raw returns, for each country using two multivariate GARCH (1,1) models: one based on equations (6a-c) that leads to our two-factor benchmark Model I and another based on the full set of equations (6a-f). The test statistic shows substantial capturing of non-normality of equity market returns across all countries. The biggest improvement is observed after applying the first GARCH model based on equations (6a-c).

more capital constraints from shrinking global credit supply than their peers in the developed world.

We also analyze the cross-sectional properties of bond illiquidity betas. Figure 2 shows the relation between average country-level excess returns and the average conditional betas for the full sample period. We observe that most average bond illiquidity betas are negative and that there is a downward trend between these betas and mean excess returns. The plot implies that the lower in absolute terms is a country's stock market exposure to the illiquidity of U.S. Treasuries, the lower is its expected return. Not surprisingly, the set of observations with negative bond illiquidity betas (the 18 leftmost points) belongs to emerging markets or developed markets that were classified as emerging during a substantial part of our sample period, such as Greece and Portugal. The vast majority of countries with close to zero or positive bond illiquidity risk are associated with developed and therefore more liquid stock markets, which are relatively less exposed to Treasury bond illiquidity shocks, or relatively immune to the U.S. interest rate policy.

Given the wide dispersion of Treasury bond illiquidity betas across countries, we explore whether any country-specific characteristics can explain the cross-sectional differences in these betas. Table 4 reports results for country-level variables that we believe may affect bond illiquidity betas, i.e., impact the exposure of countries' equity market returns to U.S. monetary policy shifts. CORR is the average country's equity market correlation with the world market portfolio over the entire sample period. Size is the average stock market capitalization to GDP ratio from Djankov et al. (2008). LISTINGS is the number of all overseas listings traded on various world stock exchanges at the end of 1998 from Sarkissian and Schill (2004). We can think of these three variables as "market development" proxies. The more developed a country's financial market is, the lower is its exposure to U.S. monetary shocks and hence the lower are in absolute terms their bond illiquidity betas. SEG is a market segmentation proxy computed in the spirit of Bekaert et al. (2008) as the average absolute difference between a country's inverse price-to-earning ratio and that of the world market.[21] RATE is the short-term interest rate. The

[21] In Bekaert, et al. (2008), SEG is the weighted sum of local-global industry valuation differentials.

monthly price-to-earning ratios and interest rates are taken from Datastream. These two variables can be regarded as “dynamic indicators,” and are easily observable over time at any sampling frequency. The more segmented a country is from the world, or the higher is the level of its nominal interest rates, the higher is the probability for a negative reaction of its equity market to the U.S. monetary policy tightening, and hence the higher (more negative) its bond illiquidity beta is expected to be. Finally, FREEDOM is the average index of economic freedom in 1995-2006 from the Heritage Foundation,[22] and LAW is the anti-self-dealing index again from Djankov, et al. (2008). These two variables can be thought of as “investor environment” proxies. Counties with better investor protection should be associated with more developed, more liquid, and larger stock markets with lower exposure to U.S. monetary contractions, and thus should have lower bond illiquidity betas in absolute terms. This analysis will help us understand what country characteristics are associated with higher exposure to U.S. monetary policy shocks. While Thorbecke (1997) shows on U.S. data that smaller stocks are less immune from monetary policy tightening due to more difficult access to credit, the evidence at the country level is limited.

Table 5 reports the results of the regression of average conditional Treasury bond illiquidity betas across countries (46 data points) on various sets of country characteristics from Table 4. It also shows the R-squared for each regression. In all estimations, the number of foreign listings and short-term rate are taken with logs. Regression (1) includes only one regressor, CORR, which takes a positive and significant coefficient. This implies that the higher is the correlation between the local stock market and the world market the lower is its sensitivity, in absolute value, to bond illiquidity shocks. However, when we include the other two “market development” variables, i.e., SIZE and LISTINGS in Regression (2), CORR passes its sign and significance on to the number of overseas listings. Regression (3) presents the test results for the “dynamic indicators.” The coefficients on SEG and RATE are negative, as expected, but only the market segmentation proxy is marginally significant. This implies that less integrated but

[22] The index can be downloaded from the Foundation’s web site at http://www.heritage.org/research/features/index/.

open countries are generally less immune from bond illiquidity shocks. Regression (4) presents the results for the "investor environment" proxies. Consistent with our expectations, we find a positive and significant relation between FREEDOM and the bond illiquidity beta. This implies that economically, financially, and politically more sound countries have more liquid markets overall as well as possibly better access to credit. Thus they are less exposed (i.e., have lower in absolute value bond illiquidity beta) to monetary policy shocks in the U.S.

However, when we combine the "market development" variables with the "dynamic indicators" and "investor environment" proxies in Regression (5), we find that the only variable that retains its sign and statistical significance at the 5% level is the number of overseas listings. This variable also remains significant in the presence of the emerging market dummy, as shown in Regression (6). Thus, when a country is more integrated with the world market (e.g., through its foreign listing activity) the availability of credit is not restricted to its local market only leading to lower absolute value of bond illiquidity beta.

### *5.2. Asset Pricing Tests*

To further examine the cross-sectional importance of Treasury bond illiquidity for international equity market returns, we turn our attention to the results of the GMM-based asset pricing tests. We first examine the performance of our base two-factor model (Model I), the World CAPM with the Treasury bond illiquidity factor, WCAPM-LB. Table 6 shows the test results for two different instrument sets across all countries as well as separately for developed and emerging markets. Besides the point estimates of the prices of risk and their t-statistics, for each test the table also reports the degrees of freedom and the GMM J-statistic with its corresponding p-value. The estimation period is 1977-2006 for developed markets and 1987-2006 for emerging markets. In Panel A, the instrument set consists of a constant and the lagged values of the bond illiquidity shock and the world market return, while in Panel B the instrument set is as in (8). The conditional variances and covariances are obtained from the multivariate GARCH (1,1) using equations (6a-c).

Across both panels of Table 6, we observe a positive and significant price of world market portfolio risk, $\lambda_w$. Its magnitude is around 4.03 for the full sample of countries, which is in line with similar estimates in prior studies on world market integration (see, e.g., De Santis and Gerard, 1997; Bekaert, Harvey, and Lundblad, 2007). Using the estimates of $\lambda_w$ and, from the first-stage estimation, the average estimate (across all countries) of the conditional covariance between each country's equity return and the world market return, $Cov_{t-1}(r_{i,t}, r_{w,t})$, which is 0.165, we can compute the average expected equity market return for a typical country attributed to the world market risk factor, $\lambda_w Cov_{t-1}(r_{i,t}, r_{w,t})$. We find that $\lambda_w Cov_{t-1}(r_{i,t}, r_{w,t})$ is approximately equal to 8.0%. This is an economically meaningful number given that the average annual stock market excess return in our sample is 13.2% in Table 1 (1.1% times 12).

More importantly, Table 6 shows that the parameter of primary interest, the price of bond illiquidity risk, $\lambda_{LB}$, is negative, as expected, and significant at the 5% level or better in every estimation but one, both for the entire sample of countries and for the sub-samples of developed and emerging countries. Note that the decrease in statistical significance of the estimates of $\lambda_{LB}$ in emerging markets results largely from the shorter sample period. The point estimates of $\lambda_{LB}$ are between 1.12 and 1.36, in absolute terms, for the whole sample of 46 countries. We can use the values of $\lambda_{LB}$ and the average conditional covariance $Cov_{t-1}(r_{i,t}, L_{B,t})$ from the first-stage estimation to compute the average annual equity market premium attributed to bond illiquidity risk, $\lambda_{LB} Cov_{t-1}(r_{i,t}, L_{B,t})$. Our evaluation produces a range of values between 1.0% and 1.3%. This magnitude is comparable to that of the U.S. stock illiquidity premium of 1.1% per annum reported by Acharya and Pedersen (2005). We can also observe that on average the point estimates of $\lambda_{LB}$ in emerging markets are higher than in developed markets (3.7 versus 1.5). This evidence corroborates the results from the predictive regressions in Table 3, where bond illiquidity has a higher predictive impact on stock returns in emerging markets. In economic terms, the average price of risk in emerging markets (3.7 across both panels) implies that in these countries about 3.5% of annual stock market returns arises from their exposure to the Treasury bond illiquidity risk. Finally, the J-statistics indicate that we cannot reject our model in which the prices of the world market and bond illiquidity risks are set constant.

While Table 6 shows that the negative and significant price of bond illiquidity risk is a consistent outcome across different estimation settings, one cannot exclude the possibility that this result is due to other world or country-specific risk factors that are omitted from the analysis. In Table 7 we address this issue by estimating three alternative global asset pricing models: Model II, which includes an additional global factor, namely, foreign exchange rate risk, as well as Models III and IV, which consider partial market integration. In Panel A our instrument set is based on (8), while in Panel B the lagged world dividend yield in (8) is replaced with lagged world stock market illiquidity and the U.S. term spread. Due to the large number of parameters being estimated, we focus only on the full-sample results across all 46 countries.

The first column in both panels of Table 7 present the performance of Model II. The results show that $\lambda_w$ is significantly positive and $\lambda_{LB}$ is significantly negative. However, while the magnitude of $\lambda_{LB}$ is similar to that in Table 6, Model II yields a substantially lower price of world market risk (between 1.77 and 2.32). This decrease in the economic importance of the world market risk factor can be explained by the some prominence of the price of foreign exchange risk. In our tests, $\lambda_c$ is positive and significant at the 10% level in both panels. Its magnitude is equal to 4.0 on average, which translates into a foreign exchange risk premium of about 0.9% per annum.[23] The J-statistic shows no signs of model misspecification.

The second column of Table 7 shows the performance of Model III, a partial integration model that consists of two global factors (the world market return and bond illiquidity), as well as country-specific variance risk. This model thus has 48 parameters to be estimated. Similar to the earlier results, both $\lambda_w$ and $\lambda_{LB}$ are significant with positive and negative signs, respectively. However, the magnitude of $\lambda_w$ is again lower than that for the base case of Model I in Table 6. This decrease can be explained by the marginal significance of local equity market risks. The

[23] Our estimates of the price of foreign exchange risk using a currency basket are similar in sign and magnitude to the average unconditional estimates across individual currencies in Dumas and Solnik (1995), although their estimates are not significant at the 10% level.

average $\lambda_i$ in Panel B is significant at the 10% level. The average $\lambda_i$ across both panels is 0.98, which implies an average risk premium (across all 46 countries) associated with local market variance risk, $\lambda_i Var_{t-1}(r_{i,t})$, of about 7.7% per annum. The annual premiums associated with world market risk, $\lambda_w Cov_{t-1}(r_{i,t}, r_{w,t})$, and bond illiquidity risk, $\lambda_{LB} Cov_{t-1}(r_{i,t}, L_{B,t})$, are 4.5% and 0.9% per annum, respectively. Therefore, the implied average expected excess world market return based on Model III is 13.1% per annum, which is very close in magnitude to the average annual excess world market return of 13.2%.

Finally, in column three of Table 7, we test the performance of Model IV, a four-factor partial integration model that, relative to Model III, also includes a second country-specific factor (local stock market illiquidity) and, for the world market return equation, world market illiquidity. This model contains 95 parameters and is thus computationally the most intensive of the four models we consider. The results show that the price of world market portfolio risk remains positive and significant at the 5% level. Bond illiquidity risk retains its economic significance but loses some statistical power, becoming significant at the 10% level in both panels. None of the two local risks (variance and stock market illiquidity) is significant. Likewise, the price of world market illiquidity risk, $\lambda_{Lw}$, is insignificant. Importantly, the J-statistic shows that Model IV is rejected at the 1% level in Panel A and at the 5% level in Panel B. This result suggests that our methodology is sufficient to differentiate between the relative validity of various asset pricing models. Notwithstanding, even in a misspecified model bond illiquidity risk is consistently priced.

In sum, Table 7 shows that the Treasury bond illiquidity risk is important not only when the world market risk is taken into account but also in the presence of other global factors that have been shown in the past, albeit with various success, to have an impact on global equity returns. The table also confirms that other risks such as foreign exchange risk and local market variance risk may enhance the risk-return relation in many countries. We are unable to find a systematic effect of stock market illiquidity on equity returns around the world.[24] However, it does not necessarily undermine the finding of stock illiquidity effect in international markets reported in the literature. Given the commonality in illiquidity between stock and bond markets (see Chordia, Sarkar, and Subrahmanyam, 2005; and Baele, Bekaert, and Inghelbrecht, 2010) and that bond illiquidity is priced, we again confirm the illiquidity risk effect in global stock markets. This effect is manifested via bond illiquidity which, besides stock illiquidity information, also contains macro-economic shocks. It may be the case that macro-illiquidity (bond illiquidity) captures information about micro-illiquidity (stock illiquidity) for aggregate stock market returns.[25] Finally, our modeling framework allows us to differentiate the suitability of various model specifications, which helps us to identify those models that may work best in an international setting.

## 6. Robustness Tests

### *6.1. Alternative Interest Rate-Based Risk Factor*

The results in Tables 6 and 7 show the importance of Treasury bond illiquidity for the pricing of global equity returns in the presence of other risk factors. However, one concern with the above tests is that they do not include other interest rate related risk factors besides Treasury bond illiquidity. In other words, the risk factors that we control for are related to stock and foreign exchange markets but not to the bond market. This concern may become relevant if one recalls that Chen, Roll, and Ross (1986) find that the term spread is a risk factor for U.S. stock returns. We therefore test another three-factor model of full market integration, similar to our Model II, but where the currency factor is replaced with the term spread. More specifically:

[24] It is possible that in a fully conditional asset pricing setting, that is, when prices of risks are also allowed to vary over time, the importance of various risk factors may be greater that in our study. However, such analysis is beyond the scope of this paper.
[25] In unreported tests, we estimated a reduced version of Model IV without the bond illiquidity risk. In this model, the economic and statistical significance of local illiquidity in emerging markets is larger than that for Model IV (consistent with Bekaert, Harvey, and Lundblad, 2007), while the world illiquidity risk becomes significantly priced in developed markets (consistent with Lee, 2011). These results are available on request.

$$E_{t-1}\left(r_{i,t}\right)=\lambda_w\, Cov_{t-1}\left(r_{i,t}, r_{w,t}\right)+\lambda_{LB}\, Cov_{t-1}\left(r_{i,t}, L_{B,t}\right)+\lambda_{Term}\, Cov_{t-1}\left(r_{i,t}, Term_t\right), \tag{9}$$

where $Term_t$ is the term spread at time $t$ and $\lambda_{Term}$ is the price of term spread risk. In the first-stage GARCH(1,1) estimation, the term spread is modeled as an AR(1) process. In the second stage, we use GMM estimations with the same two instrument sets as in Table 7.

The test results are reported in Table 8. As before, the table shows the results across all countries as well as separately for developed and emerging markets. We can see that in spite of the inclusion of the term spread, the price of bond illiquidity risk maintains its negative sign, as well as its economic and statistical significance across all country groups and instrument sets. The price of world market risk is positive and significant across all countries and for the subset of developed markets, but it loses its significance completely in emerging markets. The loss of significance of $\lambda_w$ in emerging countries and the overall drop in its economic importance appear to be driven by the inclusion of the term spread. The price of term spread risk, $\lambda_{Term}$, is negative, but it is significant at the 5% level across all markets only in Panel B. This result is consistent with the sign of term spread risk in Chen, Roll, and Ross (1986).[26] Thus, the impact of illiquidity of U.S. Treasuries on global equity returns survives the inclusion of an alternative Treasury bond market factor. Moreover, neither of the model specifications is rejected by the J-test.

### *6.2. Alternative Treasury Bond Illiquidity Data*

In 1996, CRSP switched its data source from the Federal Reserve Bank of New York to GovPX indicative quotes. To determine whether this switch has implications on our analysis, we now estimate bond illiquidity using GovPX intraday quotes. We start our sample in 1992, the first full year with available GovPX data. The bond liquidity measure is based on intraday data from New York trading hours (7:30AM to 5:00PM EST). As before, we use trading data for off-the-run Treasury bills with up to one year to maturity. The monthly time-weighted average quoted bid-ask spread is calculated as the difference between the best bid and best ask prices. To obtain

[26] Chen, Roll, and Ross (1986) also report a negative and often significant loading on an unanticipated change in the term spread.

reliable estimates of the bid-ask spread, the following filters are used: (i) bid or offer quotes with a zero value are deleted, and (ii) a quoted bid-ask spread that is negative or more than 50 cents per $100 par value (a multiple of about 12–15 times the sample average) is deleted. Monthly estimates of illiquidity based on quoted intraday bid-ask spreads are averaged across three-, six-, and 12-month T-bills for each month. Similar to the previous analysis, we use AR(2) residuals of the estimated series to proxy for aggregate bond illiquidity, which we winsorize as before.

Table 9 shows the estimation results of Model I (WCAPM-LB), with GovPX data for the whole sample of countries as well as for subsamples of developed and emerging markets. As before, the conditional estimates of the variances and covariances are obtained from equations (6a-c) using multivariate GARCH (1,1) model.[27] Due to the short sample period, we use only one instrument set that includes a constant and the lagged values of bond illiquidity and the world market return. We find that using alternative bond illiquidity data does not qualitatively change our results. The price of world market risk is positive and significant, while the price of bond illiquidity risk is negative and significant across all estimations. While the magnitude of $\lambda_{LB}$ is different from that in Table 6 due to sample size differences, we still observe variation in pricing across country groups: $\lambda_{LB}$ is about three times higher in emerging markets than in developed markets, similar to the price difference observed in earlier estimations. Thus, Treasury bond illiquidity appears to be priced in global equity markets irrespective of the data used.

### *6.3. Alternative Stock Market Illiquidity Measure*

Our zero-return measure of stock market illiquidity (Zeros) is motivated by data limitations in emerging markets, since historical volume data are not widely available for those countries to use conventionally accepted Amihud's (2002) return volume-based illiquidity measure. Nevertheless, for the sub-periods of our sample where the data were available to compute the Amihud's measure, we repeat our GARCH(1,1) and GMM procedures. In Table 10,

[27] The shorter time-series sample that results from using the GovPX data makes the GARCH and GMM estimations less immune to initial value setting. Therefore, in the second-stage estimation we omit the first-stage observations until 1994 for developed markets and 1996 for emerging markets.

we show the test output for the global asset pricing model with illiquidity risk (Model IV) based on the new illiquidity measure. The estimation is performed for two instrument sets, as in Table 7. The most important result in this table is that while accounting for the Amihud illiquidity measure reduces the significance of the world market price of risk, the bond illiquidity price of risk is again negative and significant, with magnitude similar to that in our earlier tests.

### *6.4. Alternative VAR Estimations of Treasury Bond and Local Stock Market Illiquidity*

Finally, to assess the robustness of our panel data based results in Panel B of Table 2 we also test the relation between bond illiquidity and local stock market illiquidity using two alternatives, yet inferior methods. The first method estimates bivariate VAR(1) model with the illiquidity variable pair [$L_B$, $L_i$,] individually for each country. Then the estimates and tests statistics are averaged across 46 countries. The second method uses a panel GMM estimation of 46 bivariate VAR(1)-style models with the same variable set of [$L_B$, $L_i$,] per country but with country-specific effects.

Since standard VAR models assume the same structure for each endogeneous variable, the inclusion of country fixed effects induces their correlation with lagged regressors. To account for this potential problem, we apply forward mean-differencing to both $L_B$ and $L_i$ (as in Arellano and Bover, 1995). The instrument set has a constant and the lagged values of bond illiquidity and world stock market illiquidity. The results of these tests are in Table 11. Note that the outcomes of both alternative methods are qualitatively equivalent to those reported in Table 2, i.e., we again observe that bond illiquidity predicts local stock market illiquidity but not vice versa.

## 7. Conclusion

In this paper, we show that the illiquidity of U.S. Treasuries can be considered as a transmission channel of U.S. monetary policy shocks into the stock markets around the world.

Our results show that bond illiquidity has predictive power for both stock market returns and stock market illiquidity around the world. The Treasury bond illiquidity risk is priced in global equity markets, and it commands an economically and statistically significant premium even after controlling for other conventional factors, such as the world market return, foreign exchange rates, local stock market variance, local stock market illiquidity, and the U.S. term spread. Our findings indicate that, *ceteris paribus*, the higher is the sensitivity of an asset to an increase in the illiquidity of U.S. Treasuries due to monetary contraction, the larger is the asset's expected return.

## References:


Acharya, V., and L. Pedersen, 2005, Asset pricing with liquidity risk, Journal of Financial Economics 77, 375-410.

Acharya, V, Amihud, Y., and S. Bharath, 2009, Liquidity risk of corporate bond returns, NYU working paper.

Amihud, Y., and H. Mendelson, 1986, Asset pricing and the bid-ask spread, Journal of Financial Economics 17, 223-249.

Amihud, Y., 2002, Illiquidity and stock returns: Cross-section and time-series effects, Journal of Financial Markets 5, 31–56.

Andersen, T.G., and B. Sørensen, 1996, GMM estimation of a stochastic volatility model: A Monte Carlo study, Journal of Business and Economic Statistics 14, 328-352.

Andrews, D., and J. Monahan, 1992, An improved heteroskedasticity and autocorrelation consistent covariance matrix estimator, Econometrica 60, 953-966.

Ang, A., and G. Bekaert, 2002, International asset allocation with regime shifts, Review of Financial Studies 15, 1137-1187.

Ang, A., and J. Chen, 2002, Asymmetric correlations of equity portfolios, Journal of Financial Economics 63, 443-494.

Arellano, M., and O. Bover, 1995, Another look at the instrumental variable estimation of error component models, Journal of Econometrics 68, 29-51.

Baele, L., G. Bekaert, and K. Inghelbrecht, 2010, The determinants of stock and bond return comovements, Review of Financial Studies 23, 2374-2428.

Bailey, Warren, 1990, U.S. Money Supply Announcements and Pacific Rim Stock Markets: Evidence and Implications, Journal of International Money and Finance 9, 344-356.

Bekaert, G., and C. Harvey, 1995, Time-varying world market integration, Journal of Finance 50, 403-444.

Bekaert, G., C. Harvey, and C. Lundblad, 2007, Liquidity and expected returns: Lessons form emerging markets, Review of Financial Studies 20, 1783-1832.

Bekaert, G., C. Harvey, C. Lundblad, and S. Siegel, 2008, What segments equity markets? Working paper, Duke University.

Benston, G., and R. Hagerman, 1974, Determinants of bid-ask spreads in the over-the-counter market, Journal of Financial Economics 1, 353-364.

Bernanke, B., and A. Blinder, 1988, Credit, money, and aggregate demand, American Economic Review 78, 435-439.

Bernanke, B., and A. Blinder, 1992, The federal funds rate and the channels of monetary transmission, American Economic Review 82, 901-921.

Bernanke, B., and M. Gertler, 1995, Inside the black box: The credit channel of monetary policy transmission, Journal of Economic Perspectives 9, 27-48.

Bernarke, B., and K. Kuttner, 2005, What explains the stock market's reaction to Federal Reserve policy? Journal of Finance 60, 1221–1257.

Blundell, R., and S. Bond, 1998, Initial conditions and moment restrictions in dynamic panel data models, Journal of Econometrics 87, 115-143.

Blundell, R., and S. Bond, 2000, GMM Estimation with persistent panel data: An application to production functions, Econometric Reviews 19, 321-340.

Breen,W., L. Glosten, and R. Jagannathan, 1989, Economic significance of predictable variations in stock index returns, Journal of Finance 44, 1177-1190.

Brennan, M., and A. Subrahmanyam, 1996, Market microstructure and asset pricing: On the compensation for illiquidity in stock returns, Journal of Financial Economics 41, 441-464.

Campbell, J., 1987, Stock returns and the term structure, Journal of Financial Economics 18, 373-399.

Chan, K.C., G.A. Karolyi, and R. Stulz, 1992, Global financial markets and the risk premium on U.S. equity, Journal of Financial Economics 32, 137-167.

Chen, N.-F., R. Richard, and S. Ross, 1986, Economic forces and the stock market, Journal of Business 59, 383-403.

Chordia, T., R. Roll, and A. Subrahmanyam, 2001, Market liquidity and trading activity, Journal of Finance 56, 501-530.

Chordia, T., A. Sarkar, and A. Subrahmanyam, 2005, An empirical analysis of stock and bond market liquidity, Review of Financial Studies 18, 85-129.

De Santis, G., and B. Gerard, 1997, International asset pricing and portfolio diversification with time-varying risk, Journal of Finance 52, 1881-1912.

Djankov, S., R. La Porta, F. Lopez-de-Silanes, and A. Shleifer, 2008, The law and economics of self-dealing, Journal of Financial Economics 88, 430-465.

Dumas, B., and B. Solnik, 1995. The world price of foreign exchange risk, Journal of Finance 50, 445-479.

Engle, R., 2002, Dynamic conditional correlation - a simple class of multivariate GARCH models, Journal of Business and Economic Statistics 20, 339-350.

Engle, R., and K. Kroner, 1995, Multivariate simultaneous generalized ARCH. Econometric Theory 11, 122–150.

Errunza, V., and E. Losq, 1985, International asset pricing under mild segmentation: Theory and test, Journal of Finance 40, 105-124.

Fama, E., and G. Schwert, 1977, Asset returns and inflation, Journal of Financial Economics 5, 115-146.

Fama, E. and K. French, 1989, Business conditions and expected returns on stocks and bonds, Journal of Financial Economics 25, 23-49.

Ferson W., and S. Foerster, 1994, Finite sample properties of the generalized method of Moments in tests of conditional asset pricing models, Journal of Financial Economics 36, 29-55.

Ferson, W., and C. Harvey, 1993, The risk and predictability of international equity returns, Review of Financial Studies 6, 527-566.

Ferson, W., S. Sarkissian, and T. Simin, 2003, Spurious regressions in financial economics? Journal of Finance 58, 1393-1413.

Goyenko, R., 2006, Stock and bond pricing with liquidity Risk. Working paper, McGill University.

Goyenko, R., C. Holden, and C. Trzcinka, 2009, Do liquidity measures measure liquidity?, Journal of Financial Economics 92, 153-181.

Goyenko, R., and A. Ukhov, 2009, Stock and bond market liquidity: A long-run empirical analysis, Journal of Financial and Quantitative Analysis 44, 189-212.

Goyenko, R., A. Subrahmanyam, and A. Ukhov, 2009, The term structure of bond market liquidity, Journal of Financial and Quantitative Analysis, forthcoming.

Hameed, A., W. Kang, and S. Viswanathan, 2010, Stock market declines and liquidity, Journal of Finance 65, 257-293.

Harvey, C., 1991, The world price of covariance risk, Journal of Finance 46, 111-157.

Hasbrouck, J., 2009, Trading costs and returns for US equities: Estimating effective costs from daily data, Journal of Finance 64, 1445-1477.

Ho, T., and H. Stoll, 1983, The dynamics of dealer markets under competition, Journal of Finance 38, 1053-1074.

Jensen, G., J. Mercer, and R. Johnson, 1996, Business conditions, monetary policy and expected security returns, Journal of Financial Economics, 40, 213-238.

Karolyi, G.A., and R. Stulz, 1996, Why do markets move together? An investigation of U.S.-Japan stock return comovements, Journal of Finance 51, 951-986.

Kashyap, A. and J. Stein, 2000, What do a million observations on banks say about the transmission of monetary policy?, American Economic Review 90, 407-428.

Lee, K-H., 2011, The world price of liquidity risk, Journal of Financial Economics 99, 136-161

Lesmond, D., J. Ogden, and C. Trzcinka, 1999, A new estimate of transactions costs. Review of Financial Studies 12, 1113-1141.

Longin, F., and B. Solnik, 1995, Is the correlation in international equity returns constant: 1960-1990?, Journal of International Money and Finance, 14, 1, 3-26.

Longin, F. and B. Solnik, 2001, Extreme correlation of international equity markets, Journal of Finance 56, 651-678.

Longstaff, F., 2004, The flight-to-liquidity premium in U.S. Treasury bond prices, Journal of Business 77, 511-526.

O'Hara, M., and G. Oldfield, 1986, The microeconomics of market making, Journal of Financial and Quantitative Analysis 21, 361-376.

Pastor, L., and R. Stambaugh, 2003, Liquidity risk and expected stock returns, Journal of Political Economy 111, 642-685.

Patelis, A., 1997. Stock return predictability and the role of monetary policy, Journal of Finance 52, 1951-1972.

Rigobon, R., and B. Sack, 2004, The impact of monetary policy on asset prices, Journal of Monetary Economics 51, 1553-1575.

Romer, C., and D. Romer, 1990, New evidence on the monetary transmission mechanism, Brookings Papers on Economic Activity 1, 149-213.

Sarkissian, S., and M. Schill, 2004, The overseas listing decision: New evidence of proximity preference, Review of Financial Studies 17, 769-809.

Scruggs, J., 1998, Resolving the puzzling intertemporal relation between the market risk premium and conditional market variance: A two-factor approach, Journal of Finance 53, 575-603.

Stone, B., 1974, Systematic interest-rate risk in a two-index model of returns, Journal of Financial and Quantitative Analysis 9, 709-721.

Sweeny, R. and Warga, A., 1986, The pricing of interest rate risk: Evidence from the stock Market, Journal of Finance 41, 393-410.

Thorbecke, W., 1997, On stock market returns and monetary policy, Journal of Finance 52, 635-654.

Welch, I. and A. Goyal, 2008, A comprehensive look at the empirical performance of equity premium prediction, Review of Financial Studies, 21, 1455-1508

Wongswan, 2006, Transmission of information across international equity markets, Review of Financial Studies19, 1157-1189.

Wongswan, 2009, The response of global equity indexes to U.S. monetary policy announcements, Journal of International Money and Finance 2, 344-365.

**Table 1**

**Summary statistics**

This table presents the means, volatilities, and first-order autocorrelations of monthly excess equity returns, dividend yields, and stock market illiquidity for 23 developed and 23 emerging countries (top and bottom halves of the table). The sample period is 1977:01-2006:12. The number of observations corresponds to the number of monthly returns. The data are from Datastream and IFC. The returns are in US dollars in excess of the one-month US T-bill rate. Market illiquidity is the value-weighted average proportion of zero daily returns in a month for each market.

| | | Market return | | | Dividend yield | | | Market illiquidity | | |
|---|---|---|---|---|---|---|---|---|---|---|
| Country | Obs | Mean | σ | ρ | Mean | σ | ρ | Mean | σ | ρ |
| Australia | 360 | 0.009 | 0.070 | -0.002 | 0.334 | 0.077 | 0.948 | 0.252 | 0.092 | 0.756 |
| Austria | 360 | 0.008 | 0.062 | $0.213^{***}$ | 0.153 | 0.037 | 0.961 | 0.546 | 0.223 | 0.927 |
| Belgium | 360 | 0.008 | 0.055 | 0.081 | 0.315 | 0.124 | 0.988 | 0.311 | 0.100 | 0.715 |
| Canada | 360 | 0.006 | 0.052 | 0.039 | 0.256 | 0.091 | 0.985 | 0.205 | 0.093 | 0.851 |
| Denmark | 360 | 0.009 | 0.056 | 0.075 | 0.170 | 0.071 | 0.985 | 0.424 | 0.307 | 0.978 |
| Finland | 225 | 0.010 | 0.086 | $0.174^{***}$ | 0.202 | 0.083 | 0.966 | 0.217 | 0.120 | 0.831 |
| France | 360 | 0.010 | 0.066 | 0.074 | 0.313 | 0.117 | 0.980 | 0.217 | 0.092 | 0.556 |
| Germany | 360 | 0.006 | 0.057 | 0.011 | 0.216 | 0.077 | 0.987 | 0.162 | 0.069 | 0.554 |
| Greece | 203 | 0.013 | 0.099 | 0.075 | 0.232 | 0.097 | 0.972 | 0.142 | 0.080 | 0.473 |
| Hong Kong | 360 | 0.012 | 0.103 | $0.082^{*}$ | 0.314 | 0.108 | 0.945 | 0.304 | 0.110 | 0.649 |
| Ireland | 360 | 0.012 | 0.070 | $0.104^{**}$ | 0.346 | 0.186 | 0.986 | 0.570 | 0.226 | 0.895 |
| Italy | 360 | 0.009 | 0.075 | 0.069 | 0.217 | 0.069 | 0.969 | 0.126 | 0.068 | 0.423 |
| Japan | 360 | 0.005 | 0.064 | $0.095^{*}$ | 0.101 | 0.052 | 0.995 | 0.265 | 0.058 | 0.571 |
| Netherlands | 360 | 0.008 | 0.049 | 0.015 | 0.360 | 0.129 | 0.988 | 0.249 | 0.146 | 0.572 |
| New Zealand | 227 | 0.007 | 0.063 | -0.050 | 0.393 | 0.073 | 0.927 | 0.219 | 0.071 | 0.466 |
| Norway | 323 | 0.009 | 0.075 | 0.086 | 0.211 | 0.072 | 0.940 | 0.249 | 0.089 | 0.683 |
| Portugal | 203 | 0.005 | 0.054 | $0.138^{*}$ | 0.238 | 0.087 | 0.611 | 0.303 | 0.176 | 0.752 |
| Singapore | 360 | 0.007 | 0.085 | $0.091^{*}$ | 0.214 | 0.069 | 0.949 | 0.290 | 0.075 | 0.486 |
| Spain | 237 | 0.009 | 0.059 | 0.041 | 0.253 | 0.086 | 0.979 | 0.294 | 0.196 | 0.867 |
| Sweden | 299 | 0.011 | 0.070 | 0.079 | 0.206 | 0.062 | 0.949 | 0.212 | 0.105 | 0.783 |
| Switzerland | 360 | 0.008 | 0.051 | $0.097^{*}$ | 0.178 | 0.058 | 0.988 | 0.308 | 0.098 | 0.744 |
| United Kingdom | 360 | 0.009 | 0.065 | $0.092^{*}$ | 0.362 | 0.104 | 0.961 | 0.493 | 0.229 | 0.974 |
| United States | 360 | 0.006 | 0.044 | 0.012 | 0.267 | 0.121 | 0.994 | 0.086 | 0.047 | 0.965 |
| Argentina | 360 | 0.025 | 0.235 | 0.057 | 0.183 | 0.151 | 0.850 | 0.283 | 0.166 | 0.761 |
| Brazil | 360 | 0.017 | 0.152 | 0.027 | 0.321 | 0.228 | 0.871 | 0.527 | 0.285 | 0.913 |
| Chile | 360 | 0.016 | 0.097 | $0.203^{***}$ | 0.381 | 0.179 | 0.962 | 0.369 | 0.071 | 0.680 |
| China | 158 | 0.010 | 0.114 | -0.021 | 0.125 | 0.061 | 0.934 | 0.120 | 0.121 | -0.054 |
| Colombia | 264 | 0.019 | 0.089 | $0.365^{***}$ | 0.390 | 0.217 | 0.984 | 0.485 | 0.126 | 0.738 |
| Czech Republic | 156 | 0.009 | 0.085 | $0.200^{**}$ | 0.264 | 0.198 | 0.910 | 0.235 | 0.154 | 0.798 |
| Hungary | 156 | 0.015 | 0.104 | -0.038 | 0.128 | 0.052 | 0.863 | 0.135 | 0.097 | 0.670 |
| India | 360 | 0.009 | 0.080 | $0.127^{*}$ | 0.151 | 0.059 | 0.929 | 0.283 | 0.198 | 0.802 |
| Indonesia | 204 | 0.006 | 0.131 | $0.199^{***}$ | 0.165 | 0.093 | 0.947 | 0.359 | 0.153 | 0.803 |
| Israel | 119 | 0.007 | 0.069 | -0.026 | 0.164 | 0.066 | 0.952 | 0.148 | 0.082 | 0.256 |
| Jordan | 347 | 0.007 | 0.056 | $0.092^{*}$ | 0.26 | 0.146 | 0.919 | 0.520 | 0.110 | 0.392 |
| Korea | 360 | 0.011 | 0.108 | 0.034 | 0.162 | 0.104 | 0.925 | 0.174 | 0.083 | 0.444 |
| Malaysia | 264 | 0.005 | 0.092 | 0.091 | 0.204 | 0.080 | 0.947 | 0.309 | 0.084 | 0.598 |
| Mexico | 360 | 0.016 | 0.113 | $0.259^{***}$ | 0.187 | 0.110 | 0.942 | 0.327 | 0.131 | 0.862 |
| Pakistan | 264 | 0.012 | 0.095 | 0.060 | 0.465 | 0.262 | 0.940 | 0.310 | 0.174 | 0.715 |
| Philippines | 264 | 0.014 | 0.101 | $0.275^{***}$ | 0.132 | 0.100 | 0.970 | 0.438 | 0.178 | 0.791 |
| Poland | 156 | 0.010 | 0.116 | -0.081 | 0.131 | 0.067 | 0.910 | 0.209 | 0.112 | 0.770 |
| Russia | 119 | 0.031 | 0.164 | 0.143 | 0.095 | 0.062 | 0.898 | 0.291 | 0.131 | 0.379 |
| South Africa | 155 | 0.011 | 0.079 | 0.048 | 0.265 | 0.074 | 0.936 | 0.189 | 0.175 | 0.963 |
| Taiwan | 264 | 0.013 | 0.121 | 0.064 | 0.104 | 0.075 | 0.975 | 0.165 | 0.080 | 0.106 |
| Thailand | 360 | 0.008 | 0.100 | 0.098 | 0.275 | 0.191 | 0.943 | 0.273 | 0.066 | 0.351 |
| Turkey | 240 | 0.026 | 0.187 | 0.067 | 0.273 | 0.180 | 0.876 | 0.269 | 0.199 | 0.659 |
| Venezuela | 264 | 0.014 | 0.133 | 0.032 | 0.351 | 0.323 | 0.948 | 0.330 | 0.168 | 0.867 |
| World | 360 | 0.011 | 0.042 | 0.081 | 0.240 | 0.088 | 0.993 | 0.194 | 0.037 | 0.774 |

**Table 2**
**Determinants of stock market and Treasury bond illiquidity**
This table shows the relation between U.S. Treasury bond illiquidity and illiquidity of the world and local stock markets. The sample has 46 countries and covers a period from 1977:01 to 2006:12. Panels A and B present the results from the regression of world and local stock market illiquidity, respectively, on the lagged value of Treasury bond illiquidity and other global and local predictors. Panel C presents the results from the regression of Treasury bond illiquidity on the lagged value of world market illiquidity and other global predictors. Treasury bond illiquidity, $L_B$, is off-the-run illiquidity of Treasury bills computed from the quoted spreads available at CRSP daily Treasury files. The variables $L_w$, $r_w$, and $\sigma_w$ denote the world measures of stock market illiquidity, excess equity return, and volatility, respectively. The variables $L_i$, $r_i$, and $\sigma_i$ denote the local measures of stock market illiquidity, excess equity return, and volatility, respectively. For each market and month, illiquidity is based on the value-weighted average proportion of zero returns of all firms in a given market and month. World stock market illiquidity is the value-weighted average of countries' illiquidity. Monthly stock market volatility for each market in a given month is computed as standard deviation of daily returns in that market and month. Daily return data are from Datastream and IFC. The variables FED, TERM, MMF, and CCI denote the U.S.-based measures: change in the Fed funds rate, term spread, percentage change in the amount of funds held in money market mutual funds, and change in the consumer confidence index, respectively. The term spread is the difference in yields between the 10-year U.S. Treasury note and the one-month T-bill. The data on the amount of funds held in money market mutual funds are from the Federal Reserve Board. The consumer confidence index, which is divided by 100, is from the Conference Board. The two lags of the dependent variable are included in each regression, but their coefficients are not reported. The robust t-statistics are shown in parentheses. The estimation of $L_w$ in Panel A and $L_B$ in Panel C are performed with the Newey-West correction for six lags of the standard error. The estimation of $L_i$ in Panel B is performed with the Arellano-Bover/Blundell-Bond estimator and accounts for unobserved country and time effects. Significance at the 10%, 5% and 1% levels is denoted by *, **, and ***, respectively.

Panel A: Dependent variable: World stock market illiquidity

| | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| Constant | 0.032*** | 0.024** | 0.020* | 0.023** | 0.019* |
| | (4.04) | (2.30) | (1.84) | (2.17) | (1.78) |
| $L_{B,t-1}$ | 0.119** | 0.123** | 0.160** | 0.109* | 0.161** |
| | (2.15) | (2.18) | (2.31) | (1.65) | (2.31) |
| $r_{w,t-1}$ | | -0.001 | 0.004 | 0.001 | 0.003 |
| | | (-0.01) | (0.14) | (0.04) | (0.11) |
| $\sigma_{w,t-1}$ | | 0.606 | 0.675* | 0.711* | 0.734* |
| | | (1.42) | (1.68) | (1.74) | (1.89) |
| $FED_{t-1}$ | | | 0.254* | | 0.336 |
| | | | (1.71) | | (1.60) |
| $TERM_{t-1}$ | | | 0.018 | | 0.017 |
| | | | (1.50) | | (1.40) |
| $MMF_{t-1}$ | | | | 0.020 | |
| | | | | (0.77) | |
| $CCI_{t-1}$ | | | | 0.025 | 0.015 |
| | | | | (1.13) | (0.70) |

**Table 2 (continued)**

Panel B: Dependent variable: Local stock market illiquidity

| | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| Constant | 0.078*** | 0.094*** | 0.096*** | 0.092*** | 0.095*** |
| | (5.64) | (4.81) | (5.09) | (4.70) | (5.07) |
| $L_{B,t-1}$ | 0.674*** | 0.661*** | 0.598*** | 0.678*** | 0.625*** |
| | (4.16) | (3.58) | (3.20) | (3.57) | (3.28) |
| $r_{i,t-1}$ | | 0.012 | 0.014 | 0.009 | 0.011 |
| | | (0.76) | (0.88) | (0.60) | (0.69) |
| $\sigma_{i,t-1}$ | | 0.396* | 0.406* | 0.468** | 0.476** |
| | | (1.79) | (1.83) | (2.09) | (2.14) |
| $FED_{t-1}$ | | | 0.272* | | 0.185 |
| | | | (1.87) | | (1.22) |
| $TERM_{t-1}$ | | | -0.016 | | -0.025 |
| | | | (-0.85) | | (-1.31) |
| $MMF_{t-1}$ | | | | 0.032 | |
| | | | | (0.91) | |
| $CCI_{t-1}$ | | | | 0.063*** | 0.062*** |
| | | | | (4.59) | (4.49) |
| Country Effects | Yes | Yes | Yes | Yes | Yes |
| Time Effects | Yes | Yes | Yes | Yes | Yes |

Panel C: Dependent variable: Treasury bond market illiquidity

| | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| Constant | <-0.001 | 0.001* | 0.001 | 0.001* | 0.001 |
| | (-0.27) | (1.83) | (1.57) | (1.73) | (1.48) |
| $L_{w,t-1}$ | 0.001 | -0.002 | -0.001 | -0.003 | -0.001 |
| | (0.89) | (-0.84) | (-0.57) | (-1.02) | (-0.48) |
| $r_{w,t-1}$ | | -0.001 | <0.001 | -0.001 | <0.001 |
| | | (-0.24) | (0.12) | (-0.09) | (0.02) |
| $\sigma_{w,t-1}$ | | -0.105** | -0.080* | -0.040** | -0.063* |
| | | (-2.70) | (-1.94) | (-2.22) | (-1.70) |
| $FED_{t-1}$ | | | 0.072* | | 0.065* |
| | | | (1.82) | | (1.66) |
| $TERM_{t-1}$ | | | -0.001 | | -0.001 |
| | | | (-0.57) | | (-0.76) |
| $MMF_{t-1}$ | | | | 0.018 | |
| | | | | (1.17) | |
| $CCI_{t-1}$ | | | | 0.005* | 0.004 |
| | | | | (1.79) | (1.47) |

**Table 3**
**Predictive regressions of country equity returns**
This table presents the output of predictive regressions of country excess equity returns ($r_i$) on the lagged Treasury bond illiquidity shocks, $L_B$, as well as other lagged instruments. $L_w$ and $L_i$ are the world and country-level stock market illiquidity shocks, respectively. For each market and month, illiquidity is based on the value-weighted average proportion of zero returns of all firms in a given market and month. World stock market illiquidity is the value-weighted average of countries' illiquidity. Illiquidity shocks are the AR(2) residuals of the corresponding series. $DY_w$ and $DY_i$ are the world market and local country dividend yields, respectively. TERM is the U.S. term spread. JanD is the January dummy. Regressions in Panel B include country and year fixed effects but their coefficients are not reported. Stock market illiquidity and bond illiquidity shocks are winsorized at 1% and 99%. The t-statistics in Panel A are based on the Newey-West standard errors with six lags, while in Panel B on standard errors clustered by time. The whole sample period is 1977-2006, but it is 1987-2006 for emerging markets. Significance at the 10%, 5% and 1% levels is denoted by *, **, and ***, respectively.

Panel A: Dependent variable: World stock market return

| | Full sample period | | | Sub-periods | | |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | 1977-1986 | 1987-1996 | 1997-2006 |
| Constant | 0.005** | 0.015 | 0.008 | 0.048 | -0.040 | -0.103** |
| | (2.26) | (1.18) | (0.54) | (1.57) | (-1.20) | (-2.74) |
| $L_{B,t-1}$ | -3.054*** | -3.082*** | -3.091*** | -2.593*** | -8.678*** | -19.949* |
| | (-4.24) | (-4.24) | (-4.15) | (-3.10) | (-3.84) | (-1.94) |
| $r_{w,t-1}$ | | 0.030 | 0.028 | 0.003 | -0.016 | 0.069 |
| | | (0.56) | (0.53) | (0.03) | (-0.16) | (1.06) |
| $L_{w,t-1}$ | | -0.049 | -0.070 | -0.363* | -0.094 | 0.159 |
| | | (-0.83) | (-1.19) | (-1.87) | (-0.68) | (1.28) |
| $DY_{w,t-1}$ | | | 0.003 | 0.007 | 0.031* | 0.043*** |
| | | | (1.46) | (0.84) | (1.71) | (3.17) |
| $TERM_{t-1}$ | | | 0.022 | 0.033 | -0.032 | -0.050 |
| | | | (1.12) | (1.32) | (-0.55) | (-1.26) |
| JanD | 0.001 | 0.003 | 0.004 | 0.006 | 0.016 | 0.005 |
| | (0.17) | (0.41) | (0.52) | (0.48) | (1.16) | (0.40) |

**Table 3 (continued)**

Panel B: Dependent variable: Local stock market returns

| | Full sample period | | | Sub-periods | | |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | 1977-1986 | 1987-1996 | 1997-2006 |
| Constant | 0.028* | 0.044*** | 0.028 | 0.014 | 0.075** | 0.017 |
| | (1.74) | (2.71) | (1.60) | (0.60) | (2.54) | (0.51) |
| $L_{B,t-1}$ | -3.997*** | -3.697*** | -3.725*** | -2.484* | -8.267** | -17.515 |
| | (-2.88) | (-2.63) | (-2.77) | (-1.80) | (-2.47) | (-1.53) |
| $r_{i,t-1}$ | | 0.016 | 0.020 | -0.003 | -0.003 | 0.028 |
| | | (0.61) | (0.73) | (-0.08) | (-0.06) | (0.78) |
| $L_{i,t-1}$ | | -0.004 | -0.006 | -0.046** | -0.027 | -0.005 |
| | | (-0.42) | (-0.64) | (-2.90) | (0.95) | (-0.28) |
| $DY_{i,t-1}$ | | | 0.314*** | 0.007*** | 0.003 | 0.006*** |
| | | | (2.69) | (3.26) | (1.13) | (2.92) |
| $TERM_{t-1}$ | | | -0.041 | 0.019 | -0.199** | -0.008 |
| | | | (-0.81) | (0.38) | (-2.20) | (-0.07) |
| JanD | 0.012 | 0.009 | 0.009 | 0.005 | 0.018 | 0.012 |
| | (1.43) | (1.14) | (1.13) | (0.60) | (0.96) | (0.83) |
| Country Effects | Yes | Yes | Yes | Yes | Yes | Yes |
| Year Effects | Yes | Yes | Yes | Yes | Yes | Yes |

Panel C: Developed and Emerging markets sub-samples

| | Developed markets | | | Emerging markets | | |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Constant | 0.009 | 0.021** | 0.016 | 0.052** | 0.029 | 0.035 |
| | (1.38) | (2.34) | (1.05) | (2.40) | (1.08) | (1.13) |
| $L_{B,t-1}$ | -3.191*** | -3.233** | -3.255** | -13.580*** | -14.132** | -13.798** |
| | (-2.33) | (-2.39) | (-2.47) | (-2.64) | (-2.02) | (-1.97) |
| $r_{i,t-1}$ | | -0.001 | 0.007 | | 0.011 | 0.012 |
| | | (-0.01) | (0.26) | | (0.30) | (0.34) |
| $L_{i,t-1}$ | | -0.015*** | -0.017*** | | 0.006 | 0.003 |
| | | (-2.75) | (-2.98) | | (0.35) | (0.17) |
| $DY_{i,t-1}$ | | | 0.004*** | | | 0.003 |
| | | | (3.02) | | | (1.62) |
| $TERM_{t-1}$ | | | -0.022 | | | -0.093 |
| | | | (-0.49) | | | (-0.87) |
| JanD | 0.004 | 0.005 | 0.005 | 0.025* | 0.020 | 0.020 |
| | (0.72) | (0.89) | (0.83) | (1.67) | (1.25) | (1.28) |
| Country Effects | Yes | Yes | Yes | Yes | Yes | Yes |
| Year Effects | Yes | Yes | Yes | Yes | Yes | Yes |

**Table 4**
**Summary of macroeconomic and financial variables**
This table shows the averages of several country-level financial and macroeconomic variables. CORR is the country's equity market correlation with the world market portfolio. SIZE is the average ratio of market capitalization to GDP. LISTINGS is the number of all listings from a given country placed on foreign exchanges at the end of 1998 from Sarkissian and Schill (2004). SEG and RATE are the market segmentation proxy and short-term interest rate, respectively. SEG is computed, following Bekaert, et al. (2008), as the average absolute difference between the country's inverse price-to-earning ratio and that of the world market. The monthly price-to-earning ratios and interest rates are from Datastream. FREEDOM is the average index of economic freedom in 1995-2006 from the Heritage Foundation. LAW is the anti-self-dealing index from Djankov et al. (2008).

**Table 4 (continued)**

| Country | CORR | SIZE | LISTINGS | SEG | RATE | FREEDOM | LAW |
|---|---|---|---|---|---|---|---|
| Argentina | 0.07 | 0.58 | 19 | 8.73 | 16.5 | 66.3 | 0.34 |
| Australia | 0.58 | 1.02 | 96 | 1.45 | 7.9 | 76.7 | 0.76 |
| Austria | 0.38 | 0.16 | 12 | 1.84 | 4.2 | 70.0 | 0.21 |
| Belgium | 0.61 | 0.67 | 27 | 2.26 | 4.9 | 69.3 | 0.54 |
| Brazil | 0.26 | 0.38 | 27 | 6.29 | 24.4 | 55.8 | 0.27 |
| Canada | 0.72 | 1.06 | 266 | 1.52 | 7.2 | 73.2 | 0.64 |
| Chile | 0.20 | 0.89 | 22 | 5.08 | 0.7 | 76.0 | 0.63 |
| China | 0.16 | 0.43 | 15 | 1.41 | 6.1 | 52.6 | 0.76 |
| Colombia | 0.18 | 0.14 | 4 | 4.69 | 7.1 | 62.2 | 0.57 |
| Czech Republic | 0.39 | 0.20 | 5 | 3.93 | 7.5 | 69.6 | 0.33 |
| Denmark | 0.52 | 0.58 | 9 | 2.01 | 5.9 | 71.6 | 0.46 |
| Finland | 0.59 | 1.77 | 12 | 3.33 | 4.2 | 70.8 | 0.46 |
| France | 0.67 | 0.89 | 69 | 2.16 | 5.5 | 63.1 | 0.38 |
| Germany | 0.65 | 0.54 | 112 | 1.36 | 4.4 | 68.6 | 0.28 |
| Greece | 0.31 | 0.91 | 9 | 2.48 | 8.2 | 57.9 | 0.22 |
| Hong Kong | 0.49 | 3.61 | 19 | 2.65 | 5.1 | 90.1 | 0.96 |
| Hungary | 0.53 | 0.24 | 11 | 2.30 | 13.6 | 62.4 | 0.58 |
| India | 0.16 | 0.33 | 65 | 1.80 | 6.8 | 49.1 | 0.58 |
| Indonesia | 0.35 | 0.24 | 7 | 5.05 | 15.4 | 54.8 | 0.65 |
| Ireland | 0.62 | 0.67 | 72 | 3.51 | 7.1 | 76.6 | 0.79 |
| Israel | 0.48 | 0.53 | 65 | 2.69 | 9.2 | 63.9 | 0.73 |
| Italy | 0.49 | 0.52 | 27 | 1.06 | 5.2 | 64.6 | 0.42 |
| Japan | 0.71 | 0.69 | 206 | 3.27 | 2.2 | 70.6 | 0.50 |
| Jordan | 0.11 | 0.77 | 1 | 2.63 | 4.7 | 64.0 | 0.16 |
| Korea | 0.39 | 0.54 | 29 | 1.84 | 9.2 | 69.5 | 0.47 |
| Malaysia | 0.41 | 1.48 | 7 | 1.83 | 4.6 | 64.3 | 0.95 |
| Mexico | 0.37 | 0.21 | 30 | 4.09 | 26.2 | 61.4 | 0.17 |
| Netherlands | 0.78 | 1.31 | 105 | 3.09 | 3.2 | 73.3 | 0.20 |
| New Zealand | 0.53 | 0.40 | 22 | 2.31 | 9.9 | 81.0 | 0.95 |
| Norway | 0.59 | 0.39 | 19 | 4.42 | 7.6 | 67.4 | 0.42 |
| Pakistan | 0.12 | 0.14 | 0 | 4.61 | 9.2 | 55.9 | 0.41 |
| Philippines | 0.39 | 0.48 | 7 | 2.33 | 12.2 | 58.7 | 0.22 |
| Poland | 0.50 | 0.16 | 8 | 4.15 | 16.0 | 60.3 | 0.29 |
| Portugal | 0.54 | 0.46 | 7 | 1.77 | 4.4 | 65.0 | 0.44 |
| Russia | 0.54 | 0.33 | 6 | 7.45 | 31.0 | 49.9 | 0.44 |
| Singapore | 0.59 | 1.64 | 5 | 2.09 | 3.0 | 88.4 | 1.00 |
| South Africa | 0.60 | 1.55 | 88 | 2.91 | 13.0 | 62.8 | 0.81 |
| Spain | 0.73 | 0.79 | 24 | 2.34 | 5.6 | 66.4 | 0.37 |
| Sweden | 0.69 | 1.12 | 47 | 1.57 | 4.7 | 69.0 | 0.33 |
| Switzerland | 0.69 | 2.49 | 28 | 1.79 | 3.4 | 77.8 | 0.27 |
| Taiwan | 0.36 | 1.01 | 27 | 1.58 | 4.4 | 72.5 | 0.56 |
| Thailand | 0.36 | 0.44 | 3 | 6.28 | 7.4 | 67.8 | 0.81 |
| Turkey | 0.29 | 0.35 | 7 | 4.02 | 64.9 | 58.8 | 0.43 |
| United Kingdom | 0.70 | 1.57 | 176 | 2.30 | 8.9 | 78.4 | 0.95 |
| United States | 0.83 | 1.42 | 436 | 1.10 | 5.9 | 78.3 | 0.65 |
| Venezuela | 0.09 | 0.05 | 4 | 5.47 | 10.5 | 50.3 | 0.09 |
| Average | 0.46 | 0.79 | 49.17 | 3.11 | 9.77 | 66.89 | 0.51 |

**Table 5**
**Relation between Treasury bond illiquidity betas and macroeconomic and financial factors**

This table shows the results of regression of countries' average conditional Treasury bond illiquidity betas on the set of country-level macroeconomic and financial variables. CORR is the country's equity market correlation with the world market portfolio. SIZE is the average ratio of market capitalization to GDP. LISTINGS is the number of all listings from a given country placed on foreign exchanges at the end of 1998 from Sarkissian and Schill (2004). SEG and RATE are the market segmentation proxy and short-term interest rate, respectively. SEG is computed, following Bekaert, et al. (2008), as the average absolute difference between the country's inverse price-to-earning ratio and that of the world market. The monthly price-to-earning ratios and interest rates are from Datastream. FREEDOM is the average index of economic freedom in 1995-2006 from the Heritage Foundation. LAW is the anti-self-dealing index from Djankov et al. (2008). D(Emerging) is the emerging market dummy. The number of foreign listings and the short rate are taken with logs. The table also reports the adjusted R-squared for each regression. The t-statistics are shown in parenthesis. Significance at the 10%, 5% and 1% levels is denoted by *, **, and ***, respectively.

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| Constant | -23.33*** | -30.517*** | 1.718 | -58.993*** | -36.254 | -23.301 |
| | (-4.16) | (-5.32) | (0.27) | (-3.67) | (-1.53) | (-0.88) |
| CORR | 24.550** | -5.456 | | | -15.891 | -25.481 |
| | (2.21) | (-0.39) | | | (-1.05) | (-1.45) |
| SIZE | | 5.705* | | | 3.192 | 3.259 |
| | | (1.68) | | | (0.78) | (0.80) |
| LISTINGS | | 5.381*** | | | 5.340*** | 5.383*** |
| | | (2.58) | | | (2.57) | (2.59) |
| SEG | | | -8.431* | | -1.397 | -1.082 |
| | | | (-1.77) | | (-0.98) | (-0.75) |
| RATE | | | -5.631 | | -1.964 | -1.304 |
| | | | (-1.49) | | (-0.58) | (-0.38) |
| FREEDOM | | | | 0.780*** | 0.411 | 0.280 |
| | | | | (2.97) | (1.16) | (0.74) |
| LAW | | | | -10.161 | -13.000 | -9.910 |
| | | | | (-1.01) | (-1.35) | (-0.99) |
| D(Emerging) | | | | | | -7.514 |
| | | | | | | (-1.07) |
| Adj $R^2$ | 0.100 | 0.220 | 0.077 | 0.134 | 0.244 | 0.247 |

**Table 6**

**Tests of the World CAPM with the world market return and Treasury bond illiquidity factors**

This table shows the estimation results of the global asset pricing model with the world market portfolios return and Treasury bond illiquidity factors (Model I, WCAPM-LB) for various instruments sets. The sample period is 1977:01(1987:01)-2006:12 for developed (emerging) markets. $\lambda_w$ is the price of world market risk and $\lambda_{LB}$ is the price of bond illiquidity risk. The estimates of conditional variances and covariances are from the multivariate GARCH (1,1) model based on equations (6a-c). The instrument set consists of a constant, C, and the lagged values of the AR(2) residual of bond illiquidity, $L_B$, the world market return, $r_w$, and the world dividend yield, $DY_w$. The robust t-statistics are shown in parenthesis. The table also shows the degrees of freedom and J-statistic with its corresponding p-value (in squared brackets). Significance at the 10%, 5% and 1% levels is denoted by *, **, and ***, respectively.

Panel A: Instruments: {C, $L_B$, $r_w$}

| | All countries | Developed | Emerging |
|---|---|---|---|
| $\lambda_w$ | 4.031*** | 3.213*** | 3.682** |
| | (3.83) | (2.66) | (2.54) |
| $\lambda_{LB}$ | -1.126*** | -1.367*** | -3.321* |
| | (-2.83) | (-3.02) | (-1.69) |
| df | 139 | 70 | 70 |
| J-stat | 116.85 | 76.947 | 55.917 |
| *p*-value | [0.913] | [0.266] | [0.889] |

Panel B: Instruments: {C, $L_B$, $r_w$, $DY_w$}

| | All countries | Developed | Emerging |
|---|---|---|---|
| $\lambda_w$ | 4.026*** | 3.780*** | 3.078** |
| | (4.93) | (3.92) | (2.42) |
| $\lambda_{LB}$ | -1.358*** | -1.632*** | -4.049** |
| | (-4.32) | (-4.59) | (-2.08) |
| df | 186 | 94 | 94 |
| J-stat | 165.67 | 105.64 | 78.66 |
| *p*-value | [0.855] | [0.194] | [0.872] |

**Table 7**
**Tests of alternative global asset pricing models**

This table shows the estimation results of three global asset pricing models. The sample period is 1977:01(1987:01)-2006:12 for developed (emerging) markets. $\lambda_w$, $\lambda_{Lw}$, $\lambda_{LB}$, and $\lambda_c$ are the prices of world market risk, world market illiquidity risk, Treasury bond illiquidity risk, and currency risk, respectively. Ave $\lambda_i$ and Ave $\lambda_{Li}$ are the average prices of local market variance risk and local market illiquidity risk, respectively, both across 46 countries. The return on the currency basket deposit is calculated as the equally weighted average change in exchange rates between the U.S. dollar and four global currencies: the British Pound, Euro, Japanese Yen, and Swiss Franc. The instrument set consists of a constant, C, and the lagged values of AR(2) residual of bond illiquidity, $L_B$, world market return, $r_w$, world dividend yield, $DY_w$, AR(2) residual of global stock market illiquidity, $L_w$, and the U.S. term spread, TERM. The robust t-statistics are shown in parenthesis. The table also shows the degrees of freedom and J-statistic with its corresponding p-value (in squared brackets). Significance at the 10%, 5% and 1% levels is denoted by *, **, and ***, respectively.

Panel A: Instruments: {C, $L_B$, $r_w$, $DY_w$}

| | Model II | Model III | Model IV |
|---|---|---|---|
| $\lambda_w$ | 1.769** | 3.721*** | 3.499** |
| | (2.08) | (3.61) | (2.46) |
| $\lambda_{LB}$ | -1.051*** | -1.636*** | -1.240* |
| | (-3.39) | (-4.47) | (-1.70) |
| $\lambda_c$ | 3.948* | | |
| | (1.86) | | |
| $\lambda_{Lw}$ | | | -0.020 |
| | | | (-0.07) |
| Ave $\lambda_i$ | | 0.567 | -0.299 |
| | | (0.91) | (-0.05) |
| Ave $\lambda_{Li}$ | | | 0.623 |
| | | | (0.55) |
| df | 185 | 140 | 93 |
| J-stat | 156.48 | 130.97 | 142.15*** |
| *p*-value | [0.937] | [0.695] | [0.001] |

Panel B: Instruments: {C, $L_B$, $r_w$, $L_w$, TERM}

| | (1) | (2) | (3) |
|---|---|---|---|
| $\lambda_w$ | 2.315*** | 2.096* | 2.819** |
| | (2.81) | (1.93) | (2.00) |
| $\lambda_{LB}$ | -1.137*** | -1.736*** | -1.480* |
| | (-5.14) | (-4.08) | (-1.74) |
| $\lambda_c$ | 4.027* | | |
| | (1.95) | | |
| $\lambda_{Lw}$ | | | 0.285 |
| | | | (0.66) |
| Ave $\lambda_i$ | | 1.122* | 2.739 |
| | | (1.89) | (0.32) |
| Ave $\lambda_{Li}$ | | | -0.785 |
| | | | (-0.51) |
| df | 232 | 187 | 140 |
| J-stat | 204.31 | 173.96 | 167.84* |
| *p*-value | [0.905] | [0.743] | [0.054] |

**Table 8**

**Global asset pricing model with the world market return, Treasury bond illiquidity, and term spread factors**

This table shows the estimation results of an asset pricing model with three global factors and various instrument sets. The sample period is 1977:01(1987:01)-2006:12 for developed (emerging) markets. Here, $\lambda_w$ , $\lambda_{LB}$, and $\lambda_{Term}$ are the prices of world market risk, Treasury bond illiquidity risk, and term spread risk, respectively. The estimates of conditional variances and covariances are from the multivariate GARCH (1,1) model using equations (6a-c) and the AR(1) equation for the U.S. term spread, $TERM_t = \phi_0 + \phi_1 TERM_{t-1} + e_{TERM,t}$. The instrument set consists of a constant, C, and the lagged values of AR(2) residual of bond illiquidity, $L_B$, world market return, $r_w$, and world dividend yield, $DY_w$. The robust t-statistics are shown in parenthesis. The table also shows the degrees of freedom and J-statistic with its corresponding p-value (in squared brackets). Significance at the 10%, 5% and 1% levels is denoted by *, **, and ***, respectively.

Panel A: Instruments: {C, $L_B$, $r_w$}

| | All countries | Developed | Emerging |
|---|---|---|---|
| $\lambda_w$ | 1.585** | 2.211** | 1.083 |
| | (2.33) | (2.55) | (0.80) |
| $\lambda_{LB}$ | -1.549*** | -1.425** | -2.078** |
| | (-3.01) | (-2.50) | (-2.29) |
| $\lambda_{Term}$ | -6.364 | -3.494 | -8.675 |
| | (-1.61) | (-0.78) | (-0.65) |
| df | 138 | 69 | 69 |
| J-stat | 126.54 | 68.44 | 64.52 |
| *p*-value | [0.748] | [0.496] | [0.630] |

Panel B: Instruments: {C, $L_B$, $r_w$, $DY_w$}

| | All countries | Developed | Emerging |
|---|---|---|---|
| $\lambda_w$ | 1.553*** | 1.832** | 1.025 |
| | (2.72) | (2.50) | (0.86) |
| $\lambda_{LB}$ | -1.958*** | -1.747*** | -2.606* |
| | (-4.41) | (-3.55) | (-1.72) |
| $\lambda_{Term}$ | -7.710*** | -5.334* | -11.032 |
| | (-2.72) | (-1.71) | (-1.05) |
| df | 185 | 93 | 93 |
| J-stat | 159.29 | 89.71 | 92.39 |
| *p*-value | [0.915] | [0.577] | [0.498] |

**Table 9**

**Tests of the WCAPM-LB with Treasury bond illiquidity based on the GovPX data**

This table shows the estimation results of the two-factor global asset pricing model (WCAPM-LB) with the world market and Treasury bond illiquidity risk using GovPX data. The sample period is 1994:01-2006:12 for developed markets and 1996:01-2006:12 for emerging. $\lambda_w$ is the price of world market risk, and $\lambda_{LB(GovPX)}$ is the price of Treasury bond illiquidity risk. The estimates of conditional variances and covariances are from the multivariate GARCH (1,1) using equations (6a-c). The instrument set consists of a constant, C, and the lagged values of AR(2) residual of bond illiquidity, $L_B$, and world market return, $r_w$. The robust t-statistics are shown in parenthesis. The table also shows the degrees of freedom and J-statistic with its corresponding p-value (in squared brackets). Significance at the 10%, 5% and 1% levels is denoted by *, **, and ***, respectively.

| | All countries | Developed | Emerging |
|---|---|---|---|
| $\lambda_w$ | 2.457*** | 3.227*** | 3.309** |
| | (5.98) | (3.38) | (2.52) |
| $\lambda_{LB(GovPX)}$ | -0.404*** | -0.377** | -1.193*** |
| | (-4.52) | (-2.00) | (-3.64) |
| df | 139 | 70 | 70 |
| J-stat | 127.38 | 72.95 | 64.94 |
| *p*-value | [0.750] | [0.381] | [0.681] |

**Table 10**

**Tests of asset pricing models with Amihud illiquidity measure**

This table shows the result of the GMM estimation of Model IV for two instrument sets. The estimation period is 1977:01(1987:01)-2006:12 for developed (emerging) markets. $\lambda_w$, $\lambda_{Lw}$, and $\lambda_{LB}$ are the prices of world market risk, world market illiquidity risk, and Treasury bond illiquidity risk, respectively. Ave $\lambda_i$ and Ave $\lambda_{Li}$ are the average prices of local market variance risk and local market illiquidity risk, respectively. Stock market illiquidity is based on the Amihud measure. The instrument set consists of a constant, C, and the lagged values of AR(2) residual of bond illiquidity, $L_B$, world market return, $r_w$, world dividend yield, $DY_w$, AR(2) residual of the global stock market illiquidity, $L_w$, and the U.S. term spread, TERM. The robust t-statistics are shown in parenthesis. The table also shows degreed of freedom, as well as the value of the J-test with the corresponding p-value (in squared brackets). Significance at the 10%, 5% and 1% levels is denoted by *, **, and ***, respectively.

| | Instruments: {C, $L_B$, $r_w$, $DY_w$} | Instruments: {C, $L_B$, $r_w$, $L_w$, TERM} |
|---|---|---|
| $\lambda_w$ | 4.527 | 3.367 |
| | (1.09) | (1.23) |
| $\lambda_{LB}$ | -2.082** | -1.670*** |
| | (-2.21) | (-2.56) |
| $\lambda_{Lw}$ | 0.065 | 0.010 |
| | (0.49) | (0.11) |
| Ave $\lambda_i$ | -0.279 | -0.068 |
| | (-1.07) | (-0.83) |
| Ave $\lambda_{Li}$ | -0.123 | 0.089 |
| | (-0.28) | (0.33) |
| df | 93 | 140 |
| J-stat | 77.85 | 146.67 |
| *p*-value | [0.870] | [0.333] |

**Table 11**

**Alternative estimations of the relation between the illiquidity of Treasuries and local stock markets**

This table reports the estimates of lagged coefficients and their t-statistics on Treasury bond illiquidity, $L_B$, and local stock market illiquidity, $L_i$, as predictors of local stock market illiquidity and bond illiquidity, respectively, obtained using two different methodologies. The first method estimates bivariate VAR(1) model with the variable set [$L_B$, $L_i$] for each country. The reported statistics are averaged across 46 countries. The second method uses a panel GMM estimation of 46 bivariate VAR(1) models with the variable set [$L_B$, $L_i$] per country and with country-specific intercepts. The instrument set consists of a constant and the lagged values of bond illiquidity and world stock market illiquidity.

| | Country-by-country estimation via VAR | | Whole panel estimation via GMM | |
|---|---|---|---|---|
| | $L_{i,t}$ | $L_{i,t-1}$ | $L_{i,t}$ | $L_{i,t-1}$ |
| $L_{B,t-1}$ | 3.884** (2.49) | | 10.549** (1.98) | |
| $L_{B,t}$ | | <0.001 (0.48) | | 0.001 (0.69) |

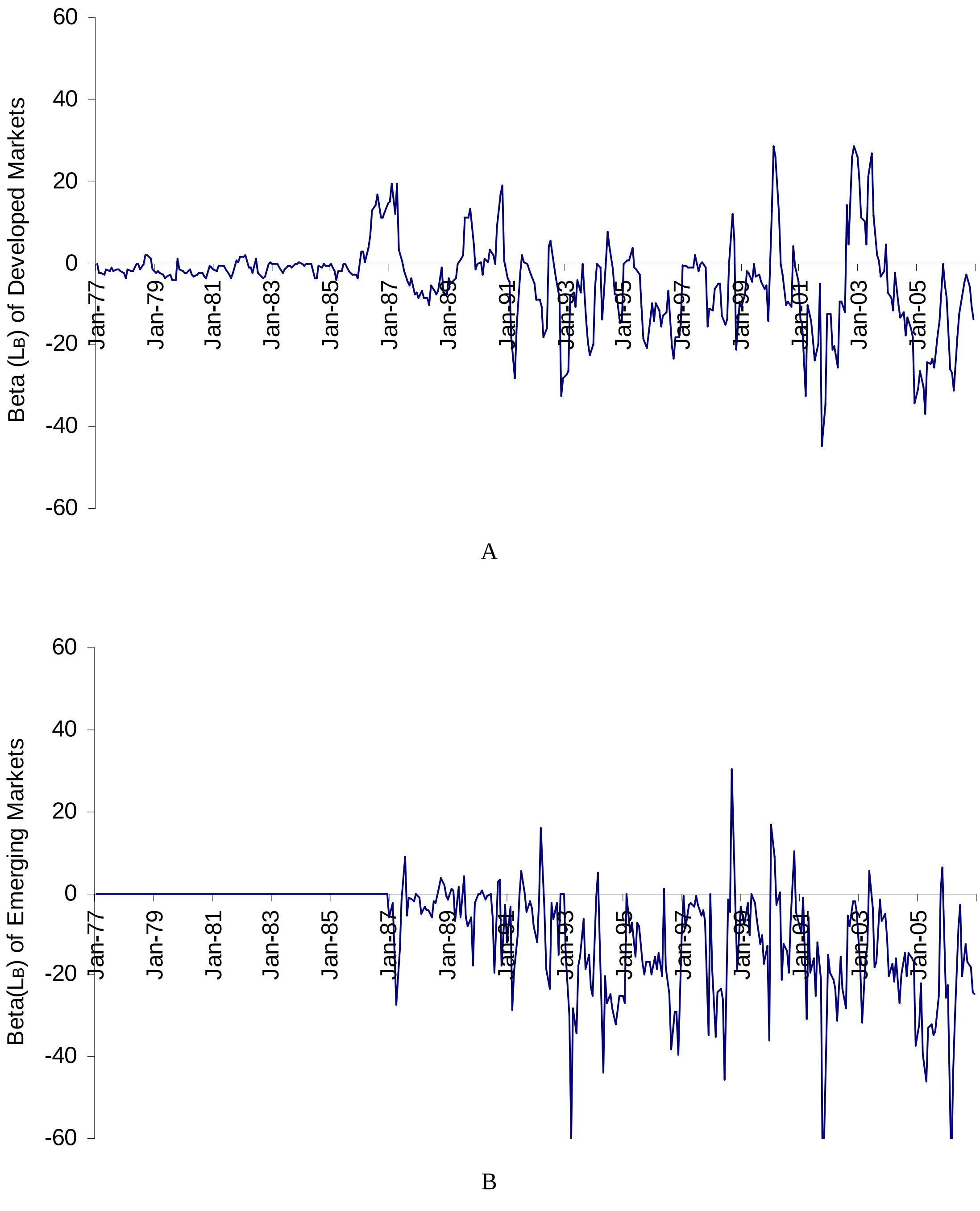


**Figure 1. Conditional Treasury bond illiquidity beta.** The figure shows average conditional Treasury bond illiquidity betas for 23 developed markets (Plot A) and 23 emerging markets (Plot B). The sample period is 1977:01(1987:01)-2006:12 for developed (emerging) markets. The conditional beta in each market is the ratio of the conditional covariance of country's excess returns with bond illiquidity risk over the conditional variance of bond illiquidity.

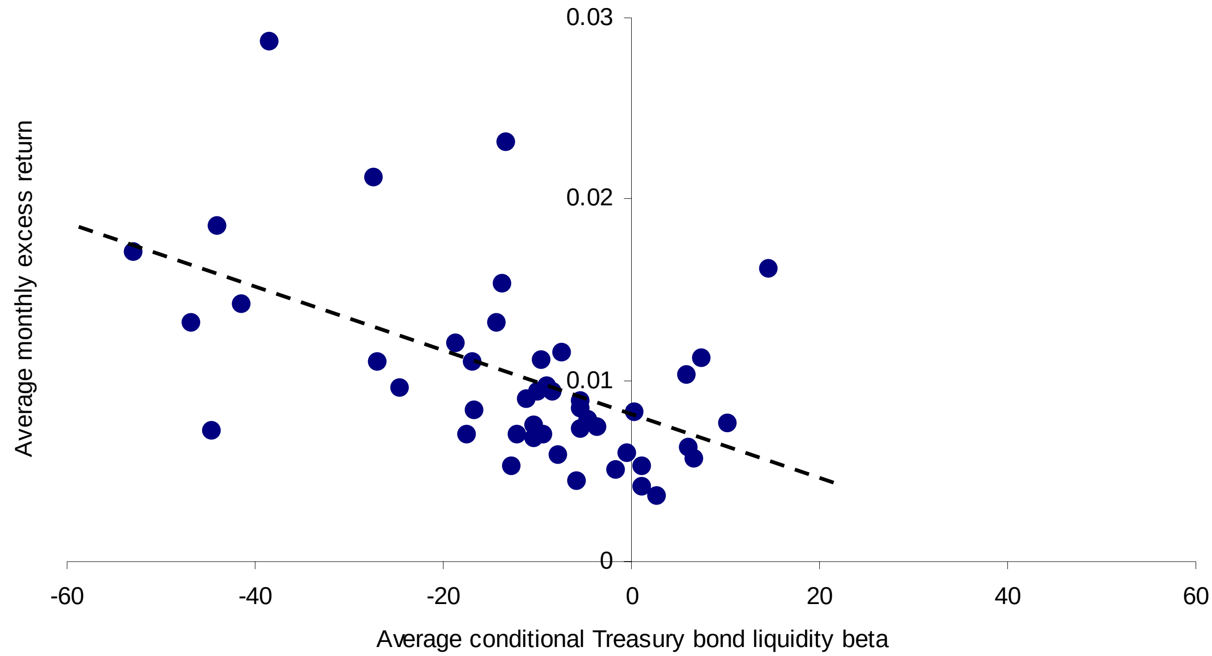


**Figure 2. Country equity returns and Treasury bond illiquidity betas.** The plot shows the relation between the mean country excess equity returns and their respective average conditional Treasury bond illiquidity betas. The sample period is 1977:01(1987:01)-2006:12 for developed (emerging) markets. The conditional beta in each market is the ratio of conditional covariance of country's excess returns with bond illiquidity risk over conditional variance of bond illiquidity. Each conditional beta is averaged over the respective sample period. The regression line is shown with a dashed line.