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Transparency, Liquidity, and Valuation: International Evidence on When Transparency Matters Most

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ABSTRACT

We examine the relation between firm-level transparency, stock market liquidity, and valuation across countries, focusing on whether the relation varies with a firm's characteristics and economic environment. We document lower transaction costs and greater liquidity (as measured by lower bid-ask spreads and fewer zero-return days) for firms with greater transparency (as measured by less evidence of earnings management, better accounting standards, higher quality auditors, more analyst following, and more accurate analyst forecasts). The relation between transparency and liquidity is more pronounced in periods of high volatility, when investor protection, disclosure requirements, and media penetration are poor, and when ownership is more concentrated, suggesting that firm-level transparency matters more

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when overall investor uncertainty is greater. Increased liquidity is associated with lower implied cost of capital and with higher valuation as measured by Tobin's *Q*. Finally, a mediation analysis suggests that liquidity is a significant channel through which transparency affects firm valuation and equity cost of capital.

1. Introduction

A substantial literature establishes a link between transparency and stock market liquidity (see, e.g., Glosten and Milgrom [1985], Kyle [1985], Welker [1995], Healy, Hutton, and Palepu [1999], and Leuz and Verrecchia [2000]). Amihud and Mendelson [1986] provide theoretical and empirical evidence that higher liquidity can lower expected returns and, thus, a firm's cost of capital.[1] Our interest is in investigating the link between transparency, liquidity, cost of capital, and firm valuation across countries and, in particular, assessing the extent to which it is influenced both by institutional and firm-level factors and by time series variation in uncertainty (we review the related literature in the next section). The international setting is especially interesting because there is substantial variation in country-, firm-, and time period-level factors that permits an examination of interactions between firm-level transparency and various aspects of the firm's economic environment.

Our sample includes 97,799 firm-year observations across 46 countries over the period 1994–2007. In our first set of tests, we relate transparency to transaction costs and stock market liquidity. To measure transparency, we employ several firm-choice variables from prior cross-country research including earnings management (Fan and Wong [2002] and Leuz, Nanda, and Wysocki [2003]), auditor quality (Fan and Wong [2005]), and adoption of global accounting standards (Daske et al. [2008, 2009]).[2] We also employ two transparency variables that capture external information gathering by intermediaries: the number of analysts who cover a firm and the accuracy of analyst forecasts.[3]

To capture transaction costs and liquidity, we use two measures that are readily available for large samples of firms across many countries and that have been shown to correlate well with actual transaction costs for

[1] For a review of the liquidity literature, see Amihud and Mendelson [2008].

[2] While Barth, Landsman, and Lang [2008] suggests that global accounting standards facilitate transparent reporting, Daske et al. [2009] shows that some firms may simply "adopt a label" while other firms actually make a serious commitment to increased transparency. Accordingly, we consider only cases in which the use of international accounting standards is accompanied by other likely indicators of a serious commitment.

[3] Several papers indicate that analysts gather and aggregate information from public and private sources to assess firm value, improving overall transparency (Brennan and Subrahmanyam [1995], Lang and Lundholm [1996], Roulstone [2003], and Lang, Lins, and Miller [2004]). In addition, research such as Yu [2008] suggests that analysts discipline firm reporting choices.

trading in a firm's shares: (1) the proportion of zero-return trading days over the fiscal year, and (2) the median bid-ask spread over the fiscal year. Bid-ask spreads speak more directly to transaction costs, while zero-return days measure liquidity more directly and are available for a wider sample of firms.[4]

We find that greater transparency—as measured by each of our components of transparency—is significantly associated with lower transaction costs and higher liquidity. In particular, liquidity is higher and transaction costs are lower for firms with less evidence of earnings management, that are audited by top-tier audit firms and, consistent with Daske et al. [2009], that seriously commit to following international accounting standards (IAS). Also, liquidity is higher and transaction costs are lower when analyst following is higher and analysts' forecasts are more accurate. Further, the economic magnitudes suggest that transparency has an economically meaningful association with liquidity.[5]

Next, we examine the extent to which the importance of firm-level transparency differs across economic settings. We expect firm-level transparency to be especially important when investor demand for information is higher because particular aspects of the firm's environment create greater uncertainty. We examine three specific aspects: country-level, firm-level, and time period characteristics.

First, at the country level, we argue that firm-level transparency will matter more in countries with greater overall opacity. We split our sample based on several institutional measures and find that the relation between firm-level transparency and liquidity is particularly strong in countries where there is likely to be significant self-dealing, disclosure requirements are relatively weak, and media penetration is low.[6]

Second, we consider time period–specific factors by examining whether transparency is particularly important during times of greater investor uncertainty, focusing on periods in which recent country-level share price volatility was especially high. Our results confirm that liquidity tends to dry up during periods of high volatility, as would be expected. Specific to our research question, we find that the effect of high volatility on liquidity is substantially mitigated for firms with high levels of transparency,

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[4] Research such as Lesmond, Ogden, and Trzcinka [1999], Lesmond [2005], and Bekaert, Harvey and Lundblad [2007] suggests that a zero return day variable better captures priced liquidity than a variety of other measures.

[5] As with the prior literature such as Daske et al. [2008, 2009], we are careful not to claim causal links for our specific transparency variables but, rather, view them as indicators of greater transparency. Our primary interest is in variation in the strength of the relation between firm-level transparency and liquidity across economic environments. Section 3.5 discusses issues of endogeneity and self-selection in more detail.

[6] Daske et al. [2008] document that mandatory adoption of IFRS is associated with greater liquidity only in countries with strong institutions. Their results likely reflect greater regulatory oversight of accounting standards, while our results likely reflect greater demand for firm-specific information in countries with greater overall opacity.

suggesting that transparency is particularly important in periods of greater uncertainty.

Third, in terms of firm-level variation, we examine whether transparency matters most for firms expected to have internal governance problems (Leuz, Lins, and Warnock [2009]). Prior research such as McConnell and Servaes [1990] suggests that, while firm value is enhanced by concentrated ownership to a point, too much ownership concentration increases agency costs and reduces firm value. As a result, we expect transparency to be particularly important to liquidity for firms with high levels of ownership concentration. To assess this, we split our sample based on ownership concentration and find that the association between transparency and liquidity is substantially stronger for firms with expected governance problems, as reflected in more concentrated ownership. Finally, we examine whether country-level institutions interact with firm-level governance and find that firm-level transparency is particularly important for firms with concentrated ownership in countries with relatively weak minority investor protection.

While drawing causal inference is difficult, our results are robust to estimation using firm fixed effects, as well as to estimation using lagged transparency measures and lagged transparency with firm fixed effects, suggesting that our transparency variables explain within-firm variation in liquidity and do not simply reflect differences in characteristics across firms. Our results are also robust to simultaneous estimation of transparency and liquidity and to controls for self-selection. Further, our results remain strong when we specifically account for changes in firms' growth opportunities and financing needs. Finally, similar results hold for short estimation windows around liquidity shocks, suggesting that reverse causality from liquidity to transparency is less likely to drive the empirical results.

In our final set of tests, we consider the linkages between liquidity, cost of capital, and valuation to assess the extent to which transparency is associated with firm valuation through liquidity. We find that liquidity is negatively correlated with ex ante cost of capital (as measured using an analyst-forecast-based valuation model) and positively correlated with firm valuation (as measured by Tobin's Q). Mediation analysis provides evidence that liquidity significantly mediates the association between transparency and cost of capital, suggesting that liquidity is an important channel through which transparency is linked to cost of capital. We also find that transparency is associated with Tobin's Q both through liquidity and other channels, likely reflecting the fact that firm valuation is affected by cash flow effects (e.g., expected expropriation of assets) as well as cost of capital. Further, the economic magnitudes of the effects we document suggest that transparency may materially affect cost of capital and valuation through its effect on liquidity.[7]

[7] A reasonable question is why, if increased transparency reduces cost of capital and increases firm value, not all firms would seek to be as transparent as possible. However, the value

Our results make several potential contributions. First, our analyses highlight firm-specific channels through which transparency may be associated with liquidity and valuation. Our results indicate that oversight by reputable auditors, reductions in earnings management, and increases in analyst following and forecast accuracy, as well as accounting standards, are all incrementally associated with greater liquidity.[8]

Second, and more importantly, our tests are specifically designed to examine cross-setting variation in the importance of firm-level transparency to liquidity, focusing on differences in country-level institutions, firm-level governance, and time-period uncertainty. Our results suggest that firm-level transparency is especially important when country-level investor protection, disclosure, and media penetration are weak, firm-level governance is weak, and a country has experienced recent volatility shocks. These results are consistent with the notion that firm-level transparency matters most when other aspects of the firm's environment increase uncertainty.

Third, our results potentially shed light on the link between liquidity, cost of capital, and valuation in international settings. We find that greater transparency is associated with higher firm valuation and lower cost of capital around the world. More importantly, the mediation analysis suggests that a substantial portion of that relation is driven through the link between transparency and liquidity. Transparency can be costly for managers, both in terms of direct costs (e.g., hiring higher quality auditors and committing to follow IAS) and indirect costs (e.g., limiting their ability to expropriate assets and disclosing potentially sensitive information to competitors). Our results suggest that potential benefit from greater transparency may accrue to a firm's shareholders through increased liquidity and lower cost of capital.

In the next section, we review the related literature on transparency, liquidity, cost of capital, and valuation. In section 3, we discuss the data and methodology and present the results of our tests. Section 4 provides conclusions.[9]

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being considered here is to the firm's minority shareholders. To the extent that management or other stakeholders such as labor or large block holders can benefit from opacity through expropriation or excess perquisite consumption, minority shareholder value may be sacrificed (Fan and Wong [2002]).

[8] While not our primary focus, we also provide evidence on the link between earnings management, liquidity, and valuation. Prior U.S.-based research (e.g., Francis, LaFond, Olsson, and Schipper [2004]) suggests that accruals that smooth earnings are associated with a lower cost of capital, while international research (e.g., Leuz, Nanda, and Wysocki [2003]) suggests that earnings smoothing is associated with greater opacity. Our results help bridge that gap by suggesting that firms with innately smooth earnings enjoy greater liquidity, consistent with Francis et al. [2004], but firms with greater discretionary smoothing experience lower liquidity, consistent with Leuz, Nanda, and Wysocki [2003].

[9] In appendices, we provide computational details for our earnings smoothing measures and baseline tests of these measures as well as a subcomponent analysis that disaggregates our primary transparency measure.

2. *Discussion of Related Literature*

Our paper builds on several strands of the international literature.[10] Most closely related are studies that analyze the effects of changes in accounting standards on measures of liquidity or cost of capital. Leuz [2003] finds no evidence that the choice of U.S. GAAP versus IAS matters to the liquidity of firms trading on the German Neuer Market. Similarly, Daske [2006] finds no evidence that adoption of International Financial Reporting Standards (IFRS) matters to cost of capital for European firms. On the other hand, Leuz and Verrecchia [2000] show that a conversion to IAS or U.S. GAAP is associated with an increase in liquidity and reduction in cost of capital for German DAX firms. Daske et al. [2008] document, for a broad global sample of firms, that capital market benefits to IFRS adoption accrue only in countries where firms have incentives to be transparent and where legal enforcement is strong. Similarly, Daske et al. [2009] find that the benefits of IFRS adoption accrue only to "serious" adopters of IFRS. Taken together, these studies suggest that, while accounting standards have the potential to affect liquidity and the cost of capital, the effects are limited to contexts in which there is greater commitment to high-quality implementation.

Additional research such as Karolyi, Lee, and Van Dijk [2012], Lang and Maffett [2011], and Ng [2011] examines determinants of liquidity risk (liquidity variability and covariability) based on a variety of firm- and country-level characteristics. Collectively, these papers draw on prior work such as Acharya and Pedersen [2005] and Brunnermeier and Pedersen [2009] that emphasizes the importance of time variation in liquidity and find that the variability and covariability of liquidity tend to be higher for firms and countries with greater opacity, especially during crisis periods. However, these studies focus on liquidity risk and uncertainty and not on determinants of the overall level of liquidity.

Two studies examine transparency and liquidity at the country level. Eleswarapu and Venkataraman [2006] use exchange-listed ADRs to study the impact of macro-level institutions on liquidity across a range of countries and find that trading costs are lower when countries have better accounting standards and legal systems. However, by construction, all of the sample firms fall under the U.S. regulatory system. Bhattacharya, Desai, and Venkataraman [2003] investigate a number of country-level relations such as those between earnings attributes, turnover, and cost of capital, finding mixed results depending on the particular measures employed.

[10] In terms of the U.S.-based disclosure literature, Welker [1995] finds a negative correlation between disclosure ratings from the Association for Investment Management and Research (AIMR) and bid-ask spreads and Healy, Hutton, and Palepu [1999] document that firms with large increases in AIMR ratings experience increases in stock performance, institutional ownership, analyst following, and liquidity. On the other hand, Heflin, Shaw, and Wild [2005] document a negative correlation between AIMR scores and market depth and Botosan and Plumlee [2002] document a positive association between cost of capital and AIMR measures of greater timely disclosure.

Several studies assess determinants of liquidity for non-U.S. securities in specific settings. Amihud, Mendelson, and Uno [1999] show that, when 66 Japanese firms reduced the lot size required for trading, the price of their shares increased, while Amihud, Lauterbach, and Mendelson [2003] show a similar outcome for a small set of Israeli stocks that moved to a more liquid trading regime. Chaplinsky and Ramchand [2004] find that yields on similar non-U.S. debt issues are higher on the less liquid Rule 144A U.S. market than on the U.S. public bond market. Bekaert, Harvey, and Lundblad [2007] document that liquidity behaves like a priced risk factor in 19 emerging market countries in the sense that unexpected liquidity shocks are positively correlated with contemporaneous return shocks and negatively correlated with shocks to dividend yield.

Finally, several studies examine determinants of cost of capital in an international setting. Hail and Leuz [2006] provide evidence that countries with better legal institutions and investor protection enjoy a lower cost of capital and Hail and Leuz [2009] suggest that firms that cross list experience a reduction in cost of capital. However, these studies do not directly examine the effect of liquidity on cost of capital nor do they focus directly on the link to transparency.

3. Data, Methodology, and Results

3.1 SAMPLE CONSTRUCTION

Accounting and market data are collected from Thomson Reuter's Datastream Advance database (a collaboration of market statistics from Datastream and accounting data from WorldScope) over the 1994–2007 time period. We require firm-year observations to have the necessary income statement and balance sheet data to calculate cash flows, accruals, and operating characteristic variables and to have sufficient market data to calculate the annual percentage of zero-return days. We exclude any country with less than 50 firm-year observations. In total, our sample contains 97,799 firm-year observations from 46 countries. Table 1 reports the frequency of observations by country. An advantage of our sample is that it includes a wide range of firms and is thus not dominated by the largest, most heavily followed multinational firms. As a result, our sample contains firms for which transparency issues are potentially more pronounced and the substantial variability within the sample should increase the power of our tests. Further, the wide range of institutional settings permits us to examine cross-setting variation in the strength of the relation between transparency and liquidity.

3.2 TRANSPARENCY AND LIQUIDITY

Our first hypothesis is that increased transparency will be associated with reduced transaction costs and increased liquidity. Because transparency is inherently difficult to measure and has many potential facets, we consider several measures.

TABLE 1
Sample by Country

Country	*N*	Percent	Country	*N*	Percent
ARGENTINA	300	0.31	LUXEMBOURG	92	0.09
AUSTRALIA	2,984	3.05	MALAYSIA	4,088	4.18
AUSTRIA	498	0.51	MEXICO	673	0.69
BELGIUM	710	0.73	NETHERLANDS	1,439	1.47
BRAZIL	1,135	1.16	NEW ZEALAND	356	0.36
CANADA	4,098	4.19	NORWAY	739	0.76
CHILE	931	0.95	PAKISTAN	287	0.29
CHINA	4,545	4.65	PERU	207	0.21
COLOMBIA	175	0.18	PHILIPPINES	806	0.82
CZECH REPUBLIC	137	0.14	POLAND	187	0.19
DENMARK	1,109	1.13	PORTUGAL	476	0.49
FINLAND	826	0.84	SINGAPORE	2,337	2.39
FRANCE	4,503	4.60	SOUTH AFRICA	1,200	1.23
GERMANY	4,388	4.49	SOUTH KOREA	3,231	3.30
GREECE	1,028	1.05	SPAIN	1,048	1.07
HONG KONG	3,386	3.46	SRI LANKA	95	0.10
HUNGARY	187	0.19	SWEDEN	1,314	1.34
INDIA	1,816	1.86	SWITZERLAND	1,497	1.53
INDONESIA	1,470	1.50	TAIWAN	3,274	3.35
IRELAND	433	0.44	THAILAND	495	0.51
ISRAEL	388	0.40	TURKEY	2,565	2.62
ITALY	1,483	1.52	UNITED KINGDOM	9,460	9.67
JAPAN	25,323	25.89	VENEZUELA	80	0.08
				97,799	100.00

The table presents the number of firms (*N*) by country for the full sample which comprises all firm-year observations between 1994 and 2007 with sufficient data from the Worldscope and Datastream databases to estimate our least restrictive specification (Model 1 for Zero Returns in table 3). Following the Datastream convention, we refer to Hong Kong as a country for simplicity. Any country with less than 50 firm-year observations is excluded from the sample.

For our first transparency variable, we estimate the degree to which a firm engages in discretionary earnings management. As discussed further in appendix A, we combine two commonly used measures of earnings management: variability of net income relative to cash flow and correlation between accruals and cash flows (e.g., Leuz, Nanda, and Wysocki [2003] and Barth, Landsman, and Lang [2008]). The underlying argument is that earnings management is manifested in the use of accruals to smooth out fluctuations in underlying cash flows.

There clearly are nondiscretionary components to earnings smoothness as well. Therefore, following the discretionary accruals literature (e.g., Jones [1991]), we first regress out a set of fundamental determinants of earnings smoothness and use the resulting residuals to form our measure of discretionary earnings smoothness. Our analyses include both the portion of smooth earnings explained by the intrinsic fundamental factors (*FUND_SMTHC*) as well as the excess portion *(DIS_SMTHC)*. Our primary interest is in *DIS_SMTHC* as a measure of transparency, and we predict that greater discretionary smoothing will be indicative of greater

earnings management and associated with greater opacity. However, we expect *FUND_SMTHC* to be positively correlated with liquidity to the extent that there is less potential for asymmetric information in firms whose profits are naturally smooth. To provide further assurance that our measure captures elements of managerial discretion, we demonstrate in appendix A that *DIS_SMTHC* is positively correlated with incentives to manage earnings (concentrated ownership and high book-tax accounting conformity) and negatively correlated with impediments to earnings management (high quality auditor, strong investor protection, global accounting standards, and high analyst following).

Additional transparency variables are also likely to be important determinants of liquidity. To the extent that analysts serve as information intermediaries, their presence should increase transparency. Lang, Lins, and Miller [2004] argue that, in an international setting, analysts are likely to play a particularly important oversight and information processing role.[11] We therefore include *ANALYST*, the number of analysts forecasting current-year earnings, as our second firm-level transparency indicator.

In addition to the number of analysts following a firm, greater accuracy of their forecasts likely reflects greater transparency of the firm's information environment. Forecast accuracy captures both the information acquisition activities of analysts as well as the disclosure policies of firms (Lang and Lundholm [1996]). Following Lang and Lundholm [1996], we measure forecast accuracy after controlling for the size of the earnings surprise and bias during the period. Thus, *R_ACCURACY,* our third transparency measure, captures, for a given magnitude of earnings surprise and bias, the extent to which analysts were able to forecast earnings.[12]

The informativeness of accounting data is likely to be higher if such data are audited by an affiliate of a global accounting firm, so we include a firm-year specific indicator variable, *BIG5*, if a firm's auditor is affiliated with a Big-5 audit firm as our fourth transparency measure (Francis [2004] and Fan and Wong [2005]).[13] Because our primary data source (Datastream) maintains firm-specific auditor data for only the most current fiscal year, we collect time-series data on firm auditor from a variety of additional sources, including historical point-in-time data from Datastream and Compustat Global. Auditor descriptions from these sources are classified as "Big-5" by hand.[14]

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[11] Research in the U.S. provides mixed evidence on the effect of analysts on transparency. Chung et al. [1995] argue that increased analyst following is associated with increased information asymmetry and reduced liquidity, while Roulstone [2003] suggests that analyst following increases publicly available information and increases liquidity.

[12] Results are not sensitive to the inclusion of controls for earnings surprise and bias.

[13] Recent scandals such as Akia Holdings, Parmalat, and Satyam have raised questions about the level of oversight provided by Big-5 firms over local affiliates.

[14] While auditor and accounting standard choices likely affect transparency through their effect on earnings management, they also likely affect transparency more generally through

Accounting data may also be more informative if a firm follows IAS, particularly if the firm is from a country with relatively low-quality local accounting regulations. However, as Daske et al. [2009] highlight, it is relatively straightforward for firms to voluntarily "adopt a label" of IAS, rather than making substantive changes, without any real economic effects. Similarly, Daske et al. [2008] find that mandatory IFRS adoption results in capital market benefits only in countries where firms have "incentives to be transparent and legal enforcement is strong" (p. 1086). Accordingly, we define serious adopters ($INTGAAP\_S = 1$), our final transparency indicator, to be adopting firms that have an above-median aggregate transparency score (calculated excluding the *INTGAAP* variable) and either a) are mandated by country regulations to adopt international accounting standards, or b) voluntarily adopted international standards (see also Barth, Landsman, and Lang [2008] and Bradshaw and Miller [2008]).[15] The overall idea is that firms with large auditors, a large and accurate analyst following, and less evidence of earnings smoothing are more likely to be adopting international accounting standards in substance rather than in form only.[16]

Our models include controls for firm size as measured by the log of a firm's market value of equity (*LNMVE*), book-to-market (*BM*), whether the firm had a loss during the period (*LOSS*), and return variability (*STDRET*), as is typical in empirical tests of liquidity (Stoll [2000]).[17] We further include indicator variables for whether the stock trades in the U.S., either on an exchange (*ADR_EX)* or on the OTC or PORTAL markets (*ADR_NEX).* U.S. trading is likely to improve transparency (Lang, Lins, and Miller [2003]), but it may also draw liquidity from local markets to the extent that shares are less costly to trade in the U.S. (Baruch, Karolyi, and Lemmon [2007]).

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their effect on overall accounting quality and the supporting disclosures provided with the financial statements.

[15] Data on accounting standards are based on Worldscope data item *WC07536*. Following Daske et al. [2009], we classify observations into three mutually exclusive categories: Local GAAP, IFRS, and U.S. GAAP; our *INTGAAP* variable includes firms following either U.S GAAP or IFRS. We distinguish voluntary and mandatory IFRS adoptions using country-specific mandatory IFRS adoption dates obtained from www.IASPLUS.com. For several countries that mandated IFRS adoption during our sample period, there exist firm-year observations where Datastream codes the accounting standard followed as local GAAP subsequent to the mandatory adoption date. Because we cannot readily determine whether these observations represent Datastream coding errors or exceptions to IFRS mandatory filing requirements, we eliminate these firms from our sample. See Daske et al. [2009] for more detail on errors in Datastream's coding of the accounting standards variable.

[16] Consistent with that conclusion (and with Daske et al. [2009]), if we include a variable for the "label" adopters, the variable is most often insignificantly related to transparency.

[17] Some prior studies also control for equity share (or dollar) turnover. As discussed in section 3.6, all of our findings are robust to the inclusion of share turnover. However, because it materially reduces the size of our sample, particularly in a subset of years and countries, we exclude share turnover from our primary analyses.

Transparency is also likely to be affected by a firm's corporate governance environment. While it is difficult to measure firm-level governance precisely, one approach is to base it on concentrated ownership. McConnell and Servaes [1990], among others, suggest that concentrated ownership may improve governance up to a certain level, but that high levels of ownership concentration are detrimental because they create sufficient control to permit, for example, expropriation of assets. To the extent that managers are entrenched, they have incentives to create opacity to hide their actions, whereas managers whose incentives are aligned with shareholders would instead want to make their firms more transparent. We therefore incorporate the percentage of closely held shares in our models, but, because the McConnell and Servaes [1990] results suggest that the relation is nonmonotonic, we split this variable into high and low ownership ranges, where *%CLHLD_H* is equal to the percentage of the firm's shares that are closely held if that proportion is above the sample median value of *%CLHLD*, and zero otherwise, and *%CLHLD_L* is equal to the percentage of the firm's shares that are closely held if that proportion is below the sample median value of *%CLHLD*, and zero otherwise.[18]

Finally, for our main specifications, we include country, year, and industry fixed effects. While transparency likely differs across countries, market microstructure does as well, so country fixed effects are important. Given potential concerns about omitted variables and endogeneity, we also report results using firm fixed effects. Within-firm comparisons have the disadvantage of ignoring potentially interesting cross-firm variation, thus weakening the power of the tests, but they have the advantage of fewer potential econometric issues.

We use two measures of liquidity and transaction costs: zero return days and bid-ask spreads. As Amihud and Mendelson [2008] note, transaction costs and liquidity are related but separate concepts. From an investor's perspective, both the direct transaction costs of trading in a firm's shares as well as the ability to form and liquidate a substantial portfolio in a timely manner are potentially important determinants of the price one is willing to pay for a stock. To measure liquidity, we follow Bekaert, Harvey, and Lundblad [2007] and define the zero-return metric (*ZERORET*) as the number of zero-return trading days over the firm's fiscal year divided by the total trading days of the fiscal year. Using the zero-return measure in our setting is advantageous because stock prices are widely available and measured consistently across markets relative to other measures such as volume or bid-ask spreads. Lesmond, Ogden, and Trzcinka [1999] argue that a manifestation

[18] Interpretation of the effect of the two *%CLHLD* variables on liquidity is further complicated because they directly reflect the supply of shares available to outside investors, meaning higher levels for each of these variables will reflect less free float (Dahlquist et al. [2003], Brockman, Chung, and Yan [2008], and Leuz, Lins, and Warnock [2009]), which could result in lower liquidity.

of illiquidity will be infrequent trading reflected in days without price movements. As such, higher values correspond to greater illiquidity.[19]

We also estimate our models with the bid-ask spread (*BIDASK*) as a proxy for transaction costs using the median bid-ask spread over the fiscal year, where the bid-ask spread is calculated as (ASK−BID)/((ASK+BID)/2), and the value is included in log form in the models following prior literature (see, e.g., Daske et al. [2008]). Given that higher values of the *ZERORET* and *BIDASK* measures correspond to greater illiquidity, we predict a negative relation between these measures and our previously described transparency measures.

3.2.1. Descriptive Statistics. Table 2, panel A, provides descriptive statistics for the variables we use in our tests, as well as several intermediate variables, grouped in the order in which they appear in subsequent tables. Our sample firms have a median market value of equity (*MVE*) of $127 million and median total assets (*ASSETS*) of $269 million. Median leverage (*LEV*), measured as debt to total assets, is 0.54, and the median book-to-market ratio (*BM*) is 0.76. The median sample firm reports losses infrequently and has experienced 6.9% sales growth (*SG*) over recent years. The table also shows that 2.2% of the sample firms have a U.S. exchange–listed ADR (*ADR_EX*) while 5.2% have an ADR that is not exchange traded (*ADR_NEX*).[20]

Statistics for our core transparency-related variables are as follows. Mean and median analyst following levels (*ANALYST*) for our sample are 3.2 and one, respectively, 36.7% of the firms are audited by Big-5 affiliates (*BIG5*), and 9.7% prepare their financial statements under nonlocal GAAP (*INTGAAP*). Turning to our liquidity variables, the median sample firm has a zero return (*ZERORET*) on 21.8% of the trading days in the year and has a bid-ask spread (*BIDASK*) of 1.3%. Data requirements to calculate *BIDASK* reduce the sample size relative to the *ZERORET* metric, with this measure available for 62,900, rather than 97,799, firm-years. We discuss average cost of capital, Tobin's *Q*, and related control variables later in the paper when we present valuation tests.

[19] Bekaert, Harvey, and Lundblad [2007] apply the zero return measure in an international context and find that the measure predicts future returns and behaves like a priced risk factor. Lesmond [2005] provides evidence that zero returns are a better proxy for liquidity than is volume or several other measures in international settings, and demonstrates that more traditional measures of transaction costs such as effective bid-ask spreads computed with detailed transaction data, where available, tend to be correlated with zero return days. Ashbaugh-Skaife, Gassen, and LaFond [2006] find that a zero return metric is a summary measure of the extent to which firm-specific information is impounded in price.

[20] Our ADR indicators are based on the 1994–2005 data used in Doidge et al. [2009], which comprises an annually collected time-series data set. These data are obtained from a variety of sources, including the Bank of New York, Citibank, JP Morgan, the NYSE, Nasdaq, the OTCBB, end-of-year editions of the National Quotation Bureau's Pink Sheets, the Center for Research in Security Prices (CRSP), firms' annual reports, SEC Form 20-F filings, and Factiva searches. We supplement the Doidge et al. [2009] data for the years 2006 and 2007 following similar procedures. Information from the various data sets is manually cross-checked and verified.

Table 2, panel A, also reports statistics for our measures of country-level institutions. The median Anti-Self Dealing Index score (*ASDI*, obtained from Djankov et al. [2008]) for our sample is 0.50 and the median country-level disclosure score (*DISCLOSE*, obtained from La Porta, Lopez-de-Silanes, and Shleifer [2006]) is 0.75. Each measure has a range of possible scores from 0 to 1, where higher values represent stronger investor protection against managerial self-dealing and greater required disclosure, respectively. *MEDIA* ranks a country's media penetration based on the World Bank's World Development Indicators. Following Bushman, Piotroski, and Smith [2004], our *MEDIA* measure features data on newspaper circulation and television ownership per capita from 1994 to 2004 and we additionally include Internet connections per capita over the same time period. The median *MEDIA* ranking in our sample is 9, where lower scores correspond to better media penetration (the best possible score is one).

TABLE 2
Descriptive Statistics

Panel A: Primary Variable Descriptive Statistics

Variable	*N*	Mean	P25	Median	P75	Std
MVE	97,799	781.239	37.621	127.023	445.638	2,857.924
BM	97,799	1.123	0.423	0.759	1.346	1.437
LOSS	97,799	0.218	0.000	0.000	0.000	0.413
STDRET	97,799	0.116	0.068	0.097	0.142	0.074
ADR_EX	97,799	0.022	0.000	0.000	0.000	0.147
ADR_NEX	97,799	0.052	0.000	0.000	0.000	0.222
%CLHLD	97,799	0.412	0.035	0.389	0.664	0.340
ANALYST	97,799	3.169	0.000	1.000	4.000	5.595
BIG5	97,799	0.367	0.000	0.000	1.000	0.482
INTGAAP	97,799	0.097	0.000	0.000	0.000	0.296
SUE	41,556	0.063	0.009	0.024	0.060	0.119
BIAS	41,556	0.019	−0.005	0.000	0.011	0.090
ACCURACY	41,556	−0.032	−0.023	−0.007	−0.002	0.086
SMTH1	97,799	−0.662	−0.877	−0.517	−0.269	0.569
SMTH2	97,799	0.735	0.688	0.920	0.983	0.410
TRANS	97,799	0.490	0.390	0.481	0.583	0.130
ZERORET	97,799	0.314	0.123	0.218	0.464	0.251
BIDASK	62,900	0.026	0.006	0.013	0.031	0.035
AVG_COC	23,605	0.114	0.080	0.102	0.133	0.054
RF_RATE	23,605	0.043	0.014	0.033	0.050	0.056
Q	76,936	1.295	0.898	1.101	1.455	0.683
ASSETS	76,936	1,069.373	96.177	269.416	852.052	2,329.830
LEV	76,936	0.530	0.387	0.537	0.673	0.211
CASH_TA	76,936	0.121	0.037	0.090	0.170	0.111
SG	76,936	0.119	0.002	0.069	0.167	0.235
NIEX_TA	76,936	0.024	0.005	0.030	0.062	0.081
DIV_DUM	76,936	0.751	1.000	1.000	1.000	0.433
CAPEX_TA	76,936	0.051	0.018	0.038	0.069	0.047
ASDI	97,799	0.594	0.460	0.500	0.760	0.233
DISCLOSE	92,651	0.732	0.667	0.750	0.833	0.162
MEDIA	89,179	14.308	5.333	9.000	20.333	10.017

(*Continued*)

TABLE 2—*Continued*

Panel B: Transparency Variable Correlation Matrix

Variable	*DIS_SMTHC*	*BIG5*	*ANALYST*	*R_ACCURACY*	*INTGAAP_S*
DIS_SMTHC	.	**−0.10**	**−0.14**	**0.02**	**−0.04**
BIG5	**−0.10**	.	**0.18**	−0.01	**0.17**
ANALYST	**−0.15**	**0.21**	.	**0.09**	**0.07**
R_ACCURACY	0.00	−0.01	**0.11**	.	**0.04**
INTGAAP_S	**−0.04**	**0.17**	**0.09**	**0.06**	.

Panel A reports descriptive statistics that are based on all firm-years between 1994 and 2007 with sufficient data to estimate the least restrictive specification of the regression model in which the data item is included. All continuous, nonlogarithmic, variables are truncated at the 1st and 99th percentiles. Definitions of variables are listed below.

Definitions of the independent variables used in liquidity regressions (tables 3–7) are as follows:

MVE is equal to the market value of equity at year end, measured in U.S. dollars (millions); *BM* is equal to the book value of common equity divided by the market value of equity; *STDRET* is the annual standard deviation of monthly stock returns; *LOSS* is equal to one if net income before extraordinary items is negative, and zero otherwise; *ADR_EX* is equal to one if the firm trades on a U.S. exchange during the year, and zero otherwise; *ADR_NEX* is equal to one if the firm has an ADR but is not traded on a U.S. exchange during the year, and zero otherwise; *%CLHLD* is the average proportion of shares that are closely held as of the end of year *t*.

Definitions of the transparency variables, and those used in the calculation thereof, are as follows:

SMTH1 is the standard deviation of net income before extraordinary items scaled by average total assets divided by the standard deviation of cash flow from operations scaled by average total assets, where standard deviations are calculated using a minimum of three and maximum of five years of data; *SMTH1* is multiplied by negative one so that larger values, i.e., values closer to zero, represent smoother earnings. *SMTH2* is defined as the correlation between the cash flow from operations and total accruals, where both measures are scaled by average total assets and correlations are calculated using a minimum of three and maximum of five years of data; *SMTH2* is multiplied by negative one so that larger values, i.e., values closer to one, represent smoother earnings. *BIG5* is equal to one if the firm is audited by a "BIG5" auditing firm, and zero otherwise; *ANALYST* is equal to the average number of analysts making a forecast for year *t*'s earnings over the *SMTH* estimation period; *ACCURACY* is the absolute value of the forecast error multiplied by −1, scaled by the stock price at the end of the prior fiscal year, where the forecast error is the I/B/E/S analysts' mean annual earnings forecast less the actual earnings as reported by I/B/E/S; *BIAS* is the signed value of the forecast error, scaled by stock price at the end of the prior fiscal year, where the forecast error is the I/B/E/S analysts' mean annual earnings forecast less the actual earnings as reported by I/B/E/S; *SUE* is the absolute value of unexpected earnings, scaled by the stock price at the end of the prior year, where unexpected earnings is actual earnings less the mean analyst forecast; *INTGAAP* is equal to one if the firm reports under IFRS or U.S. GAAP during the year, and zero otherwise; *TRANS* is calculated as described in table 3.

Definitions of the liquidity variables used in liquidity regressions (tables 3–7) are as follows:

ZERORET is equal to the percent of days in the fiscal year for which the stock price does not change; *BIDASK* is equal to the median bid ask spread over the fiscal year, where the bid ask spread is equal to (ASK−BID)/((ASK+BID)/2)).

Definitions of the valuation variables used in tables 8 and 9 are as follows:

AVG_COC is the average implied cost of capital calculated as the yearly average of four commonly used empirical estimation techniques: 1) is the modified PEG ratio model by Easton [2004], 2) is the Ohlson and Juettner-Nauroth [2005] model; 3) is the Gebhardt, Lee, and Swaminathan [2001] model; 4) is the Claus and Thomas [2001] model. Following Daske et al. [2008], we use the yearly average from these four estimation techniques as our measure of implied cost of capital. *RF_RATE* is the country-specific yearly risk-free interest rate collected from Global Insight's World Overview database. *Q* is defined as total assets less book value plus market value scaled by total assets; *ASSETS* is equal to total assets measured in U.S. dollars (millions) at the end of the fiscal year; *LEV* is equal to total debt divided by total assets measured at the end of the fiscal year; *CASH_TA* is equal to cash and cash equivalents as of the end of the fiscal year scaled by average total assets; *SG* is the average sales growth over the past three to five years; *NIEX_TA* is equal to net income before extraordinary items scaled by total assets measured at the end of the fiscal year; *DIV_DUM* is an indicator variable equal to one if the firm paid a cash dividend during the fiscal year, and zero otherwise; *CAPEX_TA* is equal to total capital expenditures for the fiscal year scaled by average total assets.

Definitions of the country-level governance variables used in table 5 are as follows:

ASDI is the Anti-Self-Dealing Index developed by Djankov et al. [2008] for the country; *DISCLOSE* is the disclosure index as reported in La Porta, Lopez-de-Silanes, and Shleifer [2006] for the country; *MEDIA* is an index constructed from the World Bank's World Development Indicators, in which each country is ranked based on the average number of newspapers, Internet connections, and televisions per capita from 1994 to 2004 and lower scores correspond to better media penetration (the best possible score is one).

Panel B reports Pearson correlation coefficients (above the diagonal) and Spearman coefficients (below the diagonal) for our five transparency variables. Details for the computation of variables not discussed above are provided in table 3. Correlations that are significant at the 5% level or better are presented in bold.

3.2.2. Correlations. Table 2, panel B, presents a correlation matrix for our transparency measures, with Pearson correlation coefficients above the diagonal and Spearman coefficients below the diagonal. No correlations exceed 0.21, suggesting that multicollinearity is unlikely to be a serious issue for our tests. That said, most measures are significantly correlated (in the predicted direction), suggesting that they capture a shared underlying construct.

3.3 LIQUIDITY RESULTS

Tables 3 and 4 display the test results for our primary hypothesis using our two measures of liquidity as dependent variables. We report results separately with and without the analyst forecast accuracy variable because its inclusion substantially reduces the sample size and limits the sample to the largest firms.

Recall that larger values of our liquidity measures represent higher illiquidity. In terms of our control variables, the results suggest that larger firms (*LNMVE*) with higher book-to-market ratios (*BM*) tend to be more liquid, although the relations are not always statistically significant. As discussed earlier, disclosure versus migration issues lead to ambiguous predictions for the two ADR coefficients. The signs of the nonexchange-traded ADR coefficients in our models are mixed, suggesting that the liquidity migration and disclosure effects tend to be offsetting when cross listing is not accompanied by an increase in mandated disclosure, but the coefficients on the exchange-traded ADR variable are generally negative, suggesting that cross listing increases liquidity when accompanied by increases in mandated disclosure and oversight.

In addition, the *%CLHLD_H* variable is significantly positive, suggesting that increases in ownership when ownership concentration is especially high are associated with lower liquidity. While this could reflect a lower level of transparency for firms with relatively concentrated ownership, it may also capture differences in free float.[21] The *%CLHLD_L* variable is (in some specifications) significantly negative, suggesting that increases in ownership when ownership concentration is relatively low have a positive net effect on governance and, hence, on firm liquidity over and above any offsetting free float effect.

Finally, the coefficient on *FUND_SMTHC* is consistently negative. This result is interesting because it suggests that firms for which accruals naturally smooth earnings relative to cash flows have higher liquidity, consistent with

[21] Consistent with a transparency explanation, *%CLHLD_H* is a significant determinant of earnings smoothing, and the relation between *%CLHLD_H* and liquidity is larger for firms with poor country-level governance (low *ASDI*, *DISCLOSURE*, and *MEDIA*) than for firms with strong country-level institutions.

TABLE 3

Zero Return Days and Transparency Determinants

	(1)		(2)		(3)		(4)		(5)		(6)	
Variable	Estimate	*p*-value	Estimate	*p*-value	Estimate	*p*-value	Estimate	*p*-value	Estimate	*p*-value	Estimate	*p*-value
Intercept	0.861	0.00	0.661	0.00	0.676	0.00	0.459	0.00	0.954	0.00	0.705	0.00
LNMVE	−0.072	0.00	−0.054	0.00	−0.058	0.00	−0.036	0.00	−0.073	0.00	−0.056	0.00
BM	−0.005	0.00	−0.003	0.01	−0.002	0.12	0.002	0.35	−0.005	0.00	−0.003	0.00
LOSS	−0.014	0.00	−0.003	0.05	−0.011	0.00	−0.002	0.22	−0.015	0.00	−0.003	0.06
STDRET	−0.461	0.00	−0.307	0.00	−0.496	0.00	−0.305	0.00	−0.462	0.00	−0.307	0.00
ADR_EX	−0.007	0.43	−0.027	0.00	0.007	0.39	−0.017	0.04	−0.015	0.09	−0.027	0.00
ADR_NEX	0.009	0.11	−0.003	0.64	0.017	0.00	0.000	0.98	0.003	0.55	−0.005	0.53
%CLHLD_H	0.039	0.00	0.029	0.00	0.032	0.00	0.016	0.00	0.045	0.00	0.031	0.00
%CLHLD_L	−0.077	0.00	0.006	0.40	−0.019	0.06	0.004	0.59	−0.061	0.00	0.009	0.24
FUND_SMTHC	−0.015	0.01	−0.022	0.00	−0.014	0.01	−0.018	0.01	−0.014	0.01	−0.023	0.00
DIS_SMTHC	0.018	0.00	0.012	0.00	0.013	0.00	0.007	0.04				
BIG5	−0.005	0.05	−0.003	0.15	0.001	0.70	−0.003	0.31				
ANALYST	−0.003	0.00	−0.003	0.00	−0.001	0.00	−0.004	0.00				
R_ACCURACY					−0.088	0.00	−0.018	0.35				
INTGAAP_S	−0.045	0.00	−0.019	0.00	−0.031	0.00	−0.029	0.00				
TRANS									−0.189	0.00	−0.081	0.00
Fixed Effects	C,I,Y		F,Y		C,I,Y		F,Y		C,I,Y		F,Y	
Observations	97,795		97,799		41,556		41,556		97,799		97,799	
S.E. Clusters (#)	F (15,681)		F (15,681)		F (8,781)		F (8,781)		F (15,681)		F (15,681)	
Adj. R^2	0.54		0.83		0.54		0.84		0.54		0.83	

The sample comprises firm-year observations between 1994 and 2007. The dependent variable in all models is *ZERORET* which is calculated as the percent of days in the fiscal year for which the stock price does not change. *%CLHLD_H* is equal to the percentage of the firm's shares which are closely held if that proportion is above the sample median value of *%CLHLD*, and zero otherwise. *%CLHLD_L* is equal to the percentage of the firm's shares which are closely held if that proportion is below the sample median value of *%CLHLD*, and zero otherwise. *FUND_SMTHC* is the average of the scaled percentile rank of *FUND_SMTH1* and *FUND_SMTH2*, where *FUND_SMTH1 & 2* are the predicted values from the fundamental earnings smoothness model described in appendix A.

The transparency variables are defined as follows: *DIS_SMTHC* is the average of the scaled percentile rank of *DIS_SMTH1* and *DIS_SMTH2*, where *DIS_SMTH1 & 2* are the residual values from the fundamental earnings smoothness model described in appendix A. *R_ACCURACY* is equal to the residual value from a regression of *ACCURACY* on *SUE* and *BIAS*. *INTGAAP_S* is an indicator variable equal to one if the firm reports under IFRS or U.S. GAAP during the year and has an above-median aggregate transparency score (calculated excluding the *INTGAAP* variable). *TRANS* is defined as the average of the scaled percentile rank of the variables: *ANALYST*, *R_ACCURACY*, *INTGAAP_S*, *BIG5*, and (1−*DIS_SMTHC*). If *R_ACCURACY* is unavailable, *TRANS* captures the average percentile rank of the remaining four measures. The control variables are defined in table 2 panel A.

We include country (C), industry (I), and year (Y) fixed effects in models 1, 3, and 5, and firm (F) and year fixed effects in models 2, 4, and 6, but do not report the coefficients. Industry fixed effects are based on the firm's two-digit ICB code. The table reports OLS coefficient estimates and p-values (two-sided) based on robust standard errors clustered at the firm-level. All continuous nonlogarithmic variables are truncated at the 1st and 99th percentiles.

$$ZERORET_t = \alpha_t + \beta_1 LNMVE_t + \beta_2 BM_t + \beta_3 LOSS_t + \beta_4 STDRET_t + \beta_5 ADR\_EX_t + \beta_6 ADR\_NEX_t + \beta_7 \%CLHLD\_H_t + \beta_8 \%CLHLD\_L_t + \beta_9 FUND\_SMTHC_t + \beta_{10} DIS\_SMTHC_t + \beta_{11} BIG5_t + \beta_{12} ANALYST_t + \beta_{13} R\_ACCURACY_t + \beta_{14} INTGAAP\_S_t + \beta_{15} TRANS_t + FIXED\ EFFECTS + \varepsilon_t$$

TABLE 4

Bid-Ask Spread and Transparency Determinants

	(1)		(2)		(3)		(4)		(5)		(6)	
Variable	Estimate	*p*-value	Estimate	*p*-value	Estimate	*p*-value	Estimate	*p*-value	Estimate	*p*-value	Estimate	*p*-value
Intercept	−2.806	0.00	−2.446	0.00	−4.284	0.00	−2.821	0.00	−2.549	0.00	−2.252	0.00
LNMVE	−0.377	0.00	−0.355	0.00	−0.332	0.00	−0.272	0.00	−0.384	0.00	−0.358	0.00
BM	0.006	0.23	−0.013	0.01	0.002	0.79	0.009	0.40	0.004	0.42	−0.014	0.01
LOSS	−0.010	0.24	0.005	0.50	−0.010	0.41	0.015	0.21	−0.014	0.09	0.004	0.57
STDRET	0.680	0.00	0.503	0.00	0.611	0.00	0.607	0.00	0.685	0.00	0.514	0.00
ADR_EX	−0.043	0.15	−0.023	0.64	−0.102	0.00	0.001	0.99	−0.077	0.01	−0.018	0.72
ADR_NEX	0.024	0.29	−0.042	0.30	0.041	0.06	−0.059	0.24	0.003	0.91	−0.040	0.32
%CLHLD_H	0.156	0.00	0.111	0.00	0.165	0.00	0.090	0.00	0.174	0.00	0.119	0.00
%CLHLD_L	−0.401	0.00	−0.048	0.21	−0.122	0.01	−0.066	0.19	−0.356	0.00	−0.041	0.29
FUND_SMTHC	−0.115	0.00	−0.164	0.00	−0.173	0.00	−0.186	0.00	−0.117	0.00	−0.163	0.00
DIS_SMTHC	0.047	0.00	0.063	0.00	0.054	0.00	0.066	0.00				
BIG5	−0.048	0.00	−0.079	0.00	−0.018	0.12	−0.041	0.01				
ANALYST	−0.010	0.00	−0.006	0.00	−0.010	0.00	−0.010	0.00				
R_ACCURACY					−0.419	0.00	−0.209	0.04				
INTGAAP_S	−0.096	0.00	−0.094	0.00	0.005	0.74	−0.040	0.01				
TRANS									−0.533	0.00	−0.356	0.00
Fixed Effects	C,I,Y		F,Y		C,I,Y		F,Y		C,I,Y		F,Y	
Observations	62,900		62,900		27,923		27,923		62,900		62,900	
S.E. Clusters (#)	F (12,382)		F (12,382)		F (6,863)		F (6,863)		F (12,382)		F (12,382)	
Adj. R^2	0.74		0.87		0.73		0.86		0.74		0.87	

The sample comprises firm-year observations between 1994 and 2007. The dependent variable in all models is *LN(BIDASK)* which is the log of *BIDASK*, where *BIDASK* is equal to the median bid ask spread over the fiscal year, calculated as (ASK-BID)/((ASK+BID)/2)). The transparency variables are defined as follows: *DIS_SMTHC* is the average of the scaled percentile rank of *DIS_SMTH1* and *DIS_SMTH2*, where *DIS_SMTH1 & 2* are the residual values from the fundamental earnings smoothness model described in appendix A. *R_ACCURACY* is equal to the residual value from a regression of *ACCURACY* on *SUE* and *BIAS*. *INTGAAP_S* is an indicator variable equal to one if the firm reports under IFRS or U.S. GAAP during the year and has an above-median aggregate transparency score (calculated excluding the *INTGAAP* variable). *TRANS* is defined as the average of the scaled percentile rank of the variables: *ANALYST*, *R_ACCURACY*, *INTGAAP_S*, *BIG5*, and (1−*DIS_SMTHC*). If *R_ACCURACY* is unavailable, *TRANS* captures the average percentile rank of the remaining four measures. The control variables are defined in table 2, panel A.

We include country (C), industry (I), and year (Y) fixed effects in models 1, 3, and 5, and firm (F) and year fixed effects in models 2, 4, and 6, but do not report the coefficients. Industry fixed effects are based on the firm's two-digit ICB code. The table reports OLS coefficient estimates and p-values (two-sided) based on robust standard errors clustered at the firm-level. All continuous nonlogarithmic variables are truncated at the 1st and 99th percentiles.

$$LN(BIDASK)_t = \alpha_t + \beta_1 LNMVE_t + \beta_2 BM_t + \beta_3 LOSS_t + \beta_4 STDRET_t + \beta_5 ADR\_EX_t + \beta_6 ADR\_NEX_t + \beta_7 \%CLHLD\_H_t + \beta_8 \%CLHLD\_L_t + \beta_9 FUND\_SMTHC_t + \beta_{10} DIS\_SMTHC_t + \beta_{11} BIG5_t + \beta_{12} ANALYST_t + \beta_{13} R\_ACCURACY_t + \beta_{14} INTGAAP\_S_t + \beta_{15} TRANS_t + FIXED\ EFFECTS + \varepsilon_t$$

research, such as Francis et al. [2004], indicating that transparency may be greater for firms for which accruals naturally smooth earnings.[22]

In terms of our primary relations of interest, we find consistent results across most specifications, indicating that our transparency measures are significantly associated with increased liquidity. For both zero-return days and bid-ask spreads, we find a significantly positive coefficient on discretionary earnings smoothing, suggesting that transaction costs are higher and investors are less willing to trade in firms' shares when managers report earnings that are excessively smoothed relative to underlying cash flows. Results in appendix A suggest that managers tend to smooth earnings more when incentives to create opacity are strongest and oversight is weakest. The results in table 3 complement those results by suggesting that earnings smoothing is associated with less liquidity. Further, the fact that the coefficient on *DIS_SMTHC* is the opposite sign from the coefficient on *FUND_SMTHC* suggests that the split between discretionary and fundamental components is not arbitrary, but instead that the two measures appear to capture different aspects of smoothness.

In addition, we find a significant negative coefficient on the indicator variable for Big-5 auditor in most specifications, suggesting that a high-quality auditor is associated with greater transparency and an increased willingness by investors to transact in a firm's shares. The coefficient on *INTGAAP_S* is negative in most specifications, suggesting, consistent with Daske et al. [2009], that use of nonlocal accounting standards by firms that are likely to be serious adopters is associated with greater liquidity.[23]

In terms of the analyst variables, both are consistent with expectations. Across all specifications, the number of analysts following the firm and analyst forecast accuracy are each positively correlated with liquidity. The results suggest that the oversight and information acquisition roles of analysts increase investors' willingness to transact in the firm's shares.

Overall, the liquidity analysis is consistent with our predictions regarding transparency. It suggests that, controlling for country, industry, year and many other factors, transaction costs are lower and liquidity is higher for firms with less evidence of earnings management, a higher quality auditor, a serious commitment to international GAAP, greater analyst following, and more accurate analyst earnings forecasts. It is noteworthy that the results are most consistent with firm fixed effects, suggesting that the findings do not simply reflect fundamental firm-level differences across the

[22] We do not include *FUND_SMTHC* as a transparency measure since it is unclear whether it reflects transparency or some other component of operating or business risk (Liu and Wysocki [2008]). However, including it among the transparency measures does not change any of our conclusions and generally strengthens our results.

[23] If we exclude mandatory adoption years from the sample, the *INTGAAP_S* results become weaker. While the coefficient remains negative in most specifications, it is significantly negatively related to liquidity only in the specifications that include country, industry, and year fixed effects and exclude the *R_ACCURACY* variable.

sample. Further, the fact that each measure has an incremental effect on transparency suggests that, while the measures are unlikely to be independent (e.g., auditor quality and accounting standards influence earnings smoothing), none subsumes the others.

For parsimony in our later analyses, we combine the transparency variables to create an aggregate measure, *TRANS*, by ranking each variable, summing the percentile ranks, and taking the average.[24] If *R_ACCURACY* is not available, the measure captures the percentile rank of the four remaining variables (*DIS_SMTHC, BIG5, ANALYST*, and *INTGAAP_S*).[25] This aggregation also reflects the fact that the various transparency measures are unlikely to be independent, so the aggregated variable captures the combined effect. We report the coefficients for models using this aggregate measure in columns 5 and 6 of tables 3 and 4. When we replace the individual transparency components with *TRANS*, the coefficient on *TRANS* is highly significant in both the *ZERORET* and *LN(BIDASK)* specifications with coefficients of −0.189 and −0.533, respectively. Appendix B reports results for the disaggregated transparency components for the analyses that follow. While the disaggregated results are generally consistent with our primary findings, the individual transparency components are not always significant.

Although our *TRANS* measure is imprecise, the economic significance of our results appears to be meaningful.[26] The results in tables 3 and 4 suggest that an interquartile shift in *TRANS* (i.e., a shift from the 25th to the 75th percentile) is associated with a nearly 17% decrease in the proportion of zero return trading days and about a 10% reduction in the bid-ask spread.[27]

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[24] The decision to aggregate the variables by adding them together is somewhat arbitrary. However, because the results are consistent across measures, conclusions are not sensitive to the aggregation approach. We also conduct a factor analysis to determine whether there are multiple factors, but the results suggest a single factor and conclusions are very similar using factor weightings. In addition, we split the measures between the analyst and other variables under the assumption that the analyst variables reflect (at least in part) external monitoring while the other variables reflect firm choices, and results are consistent across both groups.

[25] Aggregate transparency is calculated in this manner to preserve sample size, which is important because we are interested in comparisons across a wide range of firms and economic environments. As discussed in section 3.6, our inferences are robust to defining *TRANS* using only the four variables available for the entire sample.

[26] We provide estimates of economic significance only to provide intuition for the magnitudes of our coefficient estimates. We do not mean to imply that a change in our transparency variables in isolation would necessarily result in a change of liquidity of the magnitude we report since the size of the effect in practice would likely depend on the interplay of factors beyond the scope of our analysis.

[27] Economic significance for zero return days is calculated as follows: from table 2 panel A, *ZERORET* has a median value of 0.218 and an interquartile shift in *TRANS* leads to an increase of 0.193 (=0.583–0.390). The coefficient on *TRANS* in table 3 column 5 is equal to −0.189 and economic significance is calculated as ((0.218 – (0.193*−0.189))/0.218) − 1. Economic significance for the bid-ask spread is calculated similarly: from table 2 panel A, *BIDASK* has a

3.4 WHEN DOES FIRM-LEVEL TRANSPARENCY MATTER MOST?

The prior section establishes a baseline association between transparency and liquidity. However, one of our primary goals is to document the extent to which the importance of firm-level transparency varies with a firm's characteristics and its environment. We expect that firm-level transparency will be more important in countries, companies, and time periods in which there is likely to be more investor demand for information because, for example, inherent uncertainty is greater. We consider three aspects of the firm's environment that are likely to increase investor demand for information: poor country-level institutions, time periods with high aggregate levels of investor uncertainty, and firm-level governance problems.

3.4.1. Transparency and Country-Level Institutions. Country-level institutions are likely to influence the extent to which firm-level transparency affects liquidity. In particular, firm-level transparency likely matters more when investor protection (*ASDI*), disclosure (*DISCLOSE*), and media penetration (*MEDIA*) are weak. Because we estimate models with country fixed effects in our initial analysis, we cannot test for country effects using country-level variables across our full sample. To execute our tests while still controlling for country fixed effects, we split our sample into countries that score below and above the median value of these three measures, respectively.[28] Given the similarity of results for both of our liquidity measures, for parsimony we report our subsequent results using a combined liquidity measure (*ILLIQ*). *ILLIQ* is equal to the average, scaled percentile rank of the available liquidity measures, *ZERORET* and *LN(BIDASK)*. If *BIDASK* is unavailable, the measure captures the percentile rank of *ZERORET*.[29]

Table 5, panel A, reports the results of these analyses. Our primary interest is in comparing the strength of the relation between *TRANS* and *ILLIQ* across the high and low *ASDI*, *DISCLOSE*, and *MEDIA* subsamples. The results are consistent with our expectations. First, in terms of controls, most relations are similar across the two subsamples. Where the coefficients differ, they are generally consistent with expectations. For example, cross listing has a greater positive effect on liquidity for firms from countries with weak investor protection, consistent with bonding being more important in those environments. Similarly, concentrated ownership is more of an issue for liquidity in environments with weaker investor protection.

In terms of our primary relation of interest, the association between transparency and illiquidity is significantly negative for countries with both

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median value of 0.013. The coefficient on *TRANS* in table 4, column 5, is −0.533 and, using the same interquartile shift in *TRANS*, economic significance is calculated as ((0.013 – (0.193* −0.533))/0.013) − 1.

[28] Results are very similar if we estimate a single regression with an interactive term for country-level institutions and transparency.

[29] As discussed in section 3.6, inferences are robust to limiting the sample to cases in which *BIDASK* is available.

TABLE 5
Illiquidity and Country-Level Governance

Panel A: Regressions of Aggregate Illiquidity on Transparency Across Low and High Country-Level Governance Regimes

	ASDI				*DISCLOSE*				*MEDIA*			
	LOW		HIGH		LOW		HIGH		LOW		HIGH	
Variable	Estimate	*p*-value	Estimate	*p*-value	Estimate	*p*-value	Estimate	*p*-value	Estimate	*p*-value	Estimate	*p*-value
Intercept	1.183	0.00	1.378	0.00	1.156	0.00	1.215	0.00	1.144	0.00	1.364	0.00
LNMVE	−0.080	0.00	−0.093	0.00	−0.081	0.00	−0.092	0.00	−0.083	0.00	−0.094	0.00
BM	−0.006	0.00	−0.001	0.65	0.004	0.00	−0.010	0.00	0.001	0.32	−0.010	0.00
LOSS	−0.012	0.00	−0.013	0.00	−0.009	0.01	−0.013	0.00	−0.007	0.01	−0.015	0.00
STDRET	−0.282	0.00	−0.266	0.00	−0.217	0.00	−0.303	0.00	−0.173	0.00	−0.440	0.00
ADR_EX	−0.042	0.00	−0.030	0.01	−0.045	0.00	−0.026	0.02	−0.044	0.00	−0.026	0.06
ADR_NEX	−0.047	0.00	0.015	0.02	−0.031	0.02	0.005	0.40	−0.029	0.00	0.042	0.00
%CLHDL_H	0.051	0.00	0.032	0.00	0.075	0.00	0.015	0.00	0.067	0.00	−0.002	0.62
%CLHDL_L	−0.088	0.00	−0.097	0.00	−0.032	0.10	−0.124	0.00	−0.052	0.00	−0.143	0.00
FUND_SMTHC	−0.015	0.12	−0.014	0.01	−0.025	0.01	−0.008	0.14	−0.047	0.00	0.017	0.01
TRANS	−0.315	0.00	−0.099	0.00	−0.231	0.00	−0.150	0.00	−0.200	0.00	−0.139	0.00
HIGH–LOW (*p*-value)	0.216 (0.00)				0.081 (0.00)				0.061 (0.03)			
S.E. Clusters (#)	C–I (589)				C–I (589)				C–I (589)			
Fixed Effects	C,I,Y		C,I,Y		C,I,Y		C,I,Y		C,I,Y		C,I,Y	
Observations	28,260		69,539		23,485		69,166		48,062		41,117	
S.E. Clusters (#)	F (4,714)		F (10,967)		F (4,940)		F (10,741)		F (8,194)		F (7,487)	
Adj R^2	0.63		0.64		0.65		0.63		0.62		0.67	

(*Continued*)

TABLE 5—*Continued*

Panel B: Regressions of Aggregate Illiquidity on Transparency by Governance Score

	GOV_SCORE							
	(0)		(1)		(2)		(3)	
Variable	Estimate	*p*-value	Estimate	*p*-value	Estimate	*p*-value	Estimate	*p*-value
Intercept	1.209	0.00	1.142	0.00	0.770	0.00	1.159	0.00
LNMVE	−0.082	0.00	−0.080	0.00	−0.083	0.00	−0.098	0.00
BM	−0.001	0.56	0.003	0.04	−0.005	0.00	−0.010	0.00
LOSS	−0.006	0.26	−0.019	0.00	−0.007	0.11	−0.015	0.00
STDRET	−0.375	0.00	−0.090	0.00	−0.159	0.00	−0.482	0.00
ADR_EX	−0.048	0.00	−0.034	0.05	−0.033	0.11	−0.030	0.06
ADR_NEX	−0.061	0.00	0.001	0.95	−0.021	0.02	0.052	0.00
%CLHLD_H	0.050	0.00	0.086	0.00	0.035	0.00	−0.018	0.00
%CLHLD_L	−0.111	0.00	−0.012	0.64	−0.073	0.00	−0.160	0.00
FUND_SMTHC	−0.023	0.11	−0.041	0.00	−0.073	0.00	0.020	0.00
TRANS	−0.317	0.00	−0.194	0.00	−0.125	0.00	−0.117	0.00
RIGHT-LEFT (*p*-value)	0.123 (0.00)		0.069 (0.08)		0.008 (0.82)			
S.E. Clusters (#)	C–I (436)		C–I (348)		C–I (201)			
Fixed Effects	C,I,Y		C,I,Y		C,I,Y		C,I,Y	
Number of Countries	18		16		8		4	

(*Continued*)

TABLE 5—*Continued*

Panel B: Regressions of Aggregate Illiquidity on Transparency by Governance Score

	GOV SCORE							
	(0)		(1)		(2)		(3)	
Variable	Estimate	*p*-value	Estimate	*p*-value	Estimate	*p*-value	Estimate	*p*-value
Observations	15,274		18,763		21,607		42,155	
S.E. Clusters (#)	F (2,530)		F (3,255)		F (3,748)		F (6,148)	
Adj R^2	0.58		0.73		0.59		0.67	

The sample comprises firm-year observations between 1994 and 2007. The dependent variable in all models is *ILLIQ*, which is a combined measure of illiquidity equal to the average scaled percentile rank of the available liquidity measures, *ZERORET* and *LN(BIDASK)*. If *BIDASK* is unavailable, *ILLIQ* captures the percentile rank of *ZERORET*. *TRANS* is defined as the average of the scaled percentile rank of the variables: *ANALYST*, *R_ACCURACY*, *INTGAAP_S*, *BIG5*, and (1−*DIS_SMTHC*). If *R_ACCURACY* is unavailable, *TRANS* captures the average percentile rank of the remaining four measures. The control variables are defined in table 2, panel A.

In panel A, regressions are run separately for high and low *ASDI, DISCLOSE*, and *MEDIA* subgroups, where *ASDI* is the Anti-Self-Dealing Index developed by Djankov et al. [2008] for the country; *DISCLOSE* is the disclosure index as reported in La Porta, Lopez-de-Silanes, and Shleifer [2006] for the country; and *MEDIA* is an index constructed from the World Bank's World Development Indicators, in which each country is ranked based on the average number of newspapers, Internet connections, and televisions per capita from 1994 to 2004 and lower scores correspond to better media penetration (the best possible score is one). *HIGH (LOW) ASDI* firms are those firms from countries with an *ASDI* score above (below) the sample median. *HIGH (LOW) DISCLOSE* firms are those firms from countries with a *DISCLOSE* score above (below) the sample median. *HIGH (LOW) MEDIA* firms are those firms from countries with a *MEDIA* rank above (below) the sample median. *HIGH–LOW* is equal to the coefficient difference on *TRANS* between the high governance and low governance partitions.

In panel B, regressions are run separately for *GOV_SCORE* subgroups. *GOV_SCORE* is an aggregate variable composed of the three country-level institutional variables in panel A where a country receives a value of one for each of the individual institutional variables for which it is above the sample median with possible scores taking integer values between 0 and 3. All other variables are calculated as described in tables 2 and 3. *RIGHT–LEFT* is equal to the coefficient difference on *TRANS* between the column on the right and the column to the immediate left.

We include country (C), industry (I), and year (Y) fixed effects in all models, but do not report the coefficients. Industry fixed effects are based on the firm's two-digit ICB code. The table reports OLS coefficient estimates and *p*-values (two-sided) based on robust standard errors clustered at the firm-level. All continuous nonlogarithmic variables are truncated at the 1st and 99th percentiles. Assessments of significance across partitions are made based on the standard errors (clustered at the country-industry level) of the coefficient on the interaction of *TRANS* and the partitioning variable in a fully interacted specification, i.e., we allow each variable, including the fixed effects, to vary by the partitioning variable.

$$ILLIQ_t = \alpha_t + \beta_1 LNMVE_t + \beta_2 BM_t + \beta_3 LOSS_t + \beta_4 STDRET_t + \beta_5 ADR\_EX_t + \beta_6 ADR\_NEX_t + \beta_7 \%CLHLD\_H_t + \beta_8 \%CLHLD\_L_t + \beta_9 FUND\_SMTHC_t + \beta_{10} TRANS_t + FIXED\ EFFECTS + \varepsilon_t$$

weak and strong country-wide institutions across all specifications. However, consistent with our hypothesis, the relation is substantially stronger for countries with low *ASDI*, *DISCLOSE*, and *MEDIA* scores.[30] Further, the results obtain consistently regardless of whether our measure captures minority investor protection (*ASDI*), opacity (*DISCLOSE*), or the amount of information available from sources external to the firm (*MEDIA*). Coefficient estimates relating transparency to liquidity are two to three times as large for firms in countries with weak investor protection, disclosure, and media penetration, consistent with the notion that, in environments in which country-wide institutions are weak, firm-level factors such as choice of auditor, accounting standards, extent of earnings smoothing, and oversight by analysts become substantially more important.[31]

Next, we attempt to capture the incremental importance of the general governance provided by a country's institutions by creating an aggregate governance measure, *GOV_SCORE*, which combines our three country-level institutional variables. *GOV_SCORE* is calculated by summing, for each country, the number of instances *ASDI*, *DISCLOSE*, or *MEDIA* is above the sample median and, as such, ranges from zero to three. In panel B of table 5, we report results across each of the four possible *GOV_SCORE* value groups. We split the sample by group to allow coefficient estimates to vary across groups.

There are two primary takeaways from this analysis. First, the number of observations is substantial in each partition, suggesting that our sample firms and countries are relatively well dispersed amongst the four *GOV_SCORE* groupings and that the three governance measures in panel A capture different institutional attributes. More importantly, the magnitude of the *TRANS* coefficient increases monotonically, moving across partitions from strongest to weakest overall governance environments. Further, the coefficients indicate economic importance as well, with a move from the best to the worst *GOV_SCORE* group being associated with a coefficient on *TRANS* that is almost three times as large in magnitude (−0.117 vs.

[30] Assessments of significance across partitions are based on the standard errors of the coefficient on the interaction of *TRANS* and the partitioning variable in a fully interacted specification, i.e. we allow each variable, including the fixed effects, to vary by the partitioning variable. We cluster the standard errors by country-industry in these fully interacted specifications to better take account of the variation in our variable of interest.

[31] In terms of economic significance, the differences across our institutional splits are substantial. Using the same calculation method discussed in footnote 27, a shift in *TRANS* from the 25th to the 75th percentile is associated with a nearly 33% (28%, 23%) decrease in zero return days for the low *ASDI* (*DISCLOSE*, *MEDIA*) subgroup and only an 8% (11%, 10%) decrease for the high *ASDI* (*DISCLOSE*, *MEDIA*) subgroup. The differences in economic significance are similar if we instead consider differences across partitions using the *BIDASK* variable. While the economic magnitudes of the differences in the *TRANS* coefficients across institutional partitions are large, given the ability of these variables to explain substantial variation in economic outcomes shown in prior work (e.g., Djankov et al. [2008]), we do not think these differences are so large as to be implausible.

−0.317). Overall, this analysis suggests that firm-level transparency is increasingly important as additional aspects of country-level governance become worse.

3.4.2. Transparency and Time Period Uncertainty. A second environmental factor that is likely to matter for transparency is time period uncertainty. For example, when there are exogenous shocks that increase uncertainty, effects are likely to be mitigated for more transparent firms. One method for capturing market uncertainty in the U.S. is to use a measure such as the Chicago Board Options Exchange Volatility Index (VIX), which reflects expected volatility of the S&P Index over the next 30 days (e.g., Dreschler [2008]). Unfortunately, similar measures are not available for all of the markets we study here. However, investors are likely to set their expectations of future volatility based on currently observed volatility. While our preceding analyses are based on annual observations because of our measures of transparency, we can measure volatility over a much shorter window. We base our measure of expected future volatility on the last 30 days of country-level return volatility.[32] As a result, the question we ask is whether, following a month of high country-level uncertainty, firm-level illiquidity is more sensitive to transparency.[33]

Table 6 reports the empirical results of three models interacting *TRANS* with lagged country-level volatility, where *HVOL* designates a time period of above-median recent volatility. Several points are worth noting. First, *HVOL* is strongly positive across all specifications, suggesting that increases in uncertainty tend to be associated with substantial reductions in liquidity. Second, the *TRANS* variable is significantly negative across all specifications, suggesting that transparency is important to liquidity, even controlling for high investor uncertainty. Most important for our research question, the *TRANS** *HVOL* coefficient is strongly negative, suggesting that transparency is particularly important to liquidity when uncertainty is high and it helps to mitigate the effect of uncertainty on liquidity.[34] In other words, opaque firms appear to suffer substantially more from volatility shocks than do more transparent firms.

The differences in specifications are also interesting. Model 1 includes controls for only industry and year fixed effects, retaining cross-country variation. Results suggest that transparency is more important for liquidity in country/time periods with high uncertainty than those with low uncertainty. Model 2 includes controls for industry, year, and country with similar

[32] An alternative approach would be to use forecasted volatility. Results from forecasting future volatility based on past volatility are very similar.

[33] Because our transparency proxy is measured at an annual level, the assumption is that transparency measured in the preceding year is approximately representative for the subsequent 12 months. Results are very consistent if contemporaneous transparency and liquidity or alternate-length windows are used.

[34] To account for variation in *HVOL*, standard errors in the table 6 analysis are two-way clustered by firm and year-month.

TABLE 6
Illiquidity and Country-Level Volatility

	(1)		(2)		(3)	
Variable	Estimate	*p*-value	Estimate	*p*-value	Estimate	*p*-value
Intercept	0.865	0.00	1.015	0.00	0.992	0.00
LNMVE	−0.034	0.00	−0.080	0.00	−0.078	0.00
BM	−0.001	0.51	−0.006	0.00	−0.003	0.00
LOSS	0.028	0.04	0.013	0.00	0.013	0.00
STDRET	0.152	0.02	−0.387	0.00	−0.778	0.00
ADR_EX	−0.104	0.00	−0.020	0.04	−0.030	0.00
ADR_NEX	−0.059	0.00	0.000	0.95	0.003	0.67
%CLHLD_H	0.046	0.00	−0.027	0.00	−0.047	0.00
%CLHLD_L	−0.067	0.00	−0.208	0.00	−0.086	0.00
FUND_SMTHC	−0.157	0.00	−0.011	0.06	0.009	0.09
HVOL	0.142	0.00	0.059	0.00	0.037	0.00
TRANS	−0.159	0.00	−0.155	0.00	−0.041	0.00
TRANS HVOL*	−0.205	0.00	−0.062	0.00	−0.033	0.00
Fixed Effects	I,Y		C,I,Y		F,Y	
Observations	518,903		518,903		518,903	
S.E. Clusters (#)	F, Y-M (10,751, 84)		F, Y-M (10,751, 84)		F, Y-M (10,751, 84)	
Adj R^2	0.26		0.62		0.83	

The sample comprises firm-month observations between 1994 and 2007. The dependent variable in all models is *ILLIQ*, which is a combined measure of illiquidity equal to the average scaled percentile rank of the available liquidity measures, *ZERORET* and *LN(BIDASK)*. If *BIDASK* is unavailable, *ILLIQ* captures the percentile rank of *ZERORET*. *TRANS* is defined as the average of the scaled percentile rank of the variables: *ANALYST*, *R_ACCURACY*, *INTGAAP_S*, *BIG5*, and (1−*DIS_SMTHC*). If *R_ACCURACY* is unavailable, *TRANS* captures the average percentile rank of the remaining four measures. *HVOL* is an indicator variable which takes the value of one if the country-level *VOLATILITY* in month $t-1$ exceeds the country-level median, and zero otherwise. *VOLATILITY* is defined as the standard deviation of monthly stock returns. The control variables are defined in table 2, panel A.

We include industry (I) and year (Y) fixed effects in model 1; country (C), industry and year fixed effects in models 2 and 4; and year and firm (F) fixed effects in model 3, but do not report the coefficients. Industry fixed effects are based on the firm's two-digit ICB code. The table reports OLS coefficient estimates and *p*-values (two-sided) based on robust standard errors two-way clustered at the firm and year-month level. All continuous nonlogarithmic variables are truncated at the 1st and 99th percentiles.

$$\begin{aligned} ILLIQ_t = {} & \alpha_t + \beta_1 LNMVE_t + \beta_2 BM_t + \beta_3 LOSS_t + \beta_4 STDRET_t + \beta_5 ADR\_EX_t + \beta_6 ADR\_NEX_t \\ & + \beta_7 \%CLHLD\_H_t + \beta_8 \%CLHLD\_L_t + \beta_9 FUND\_SMTHC_t + \beta_{10} HVOL_t + \beta_{11} TRANS_t \\ & + \beta_{12} TRANS^* HVOL_t + FIXED\ EFFECTS + \varepsilon_t \end{aligned}$$

results, suggesting that the findings do not simply reflect cross-country variation in uncertainty. Model 3 includes year and firm fixed effects, showing that, even within a firm, transparency matters more in time periods with higher uncertainty, which suggests that our results do not simply reflect other cross-firm differences. The firm-fixed-effects finding is also important because, as discussed later, it suggests that the relation we observe between transparency and liquidity is not the result of reverse causality. From the firm's perspective, the country-wide uncertainty shock is exogenous and our firm-level transparency measure is lagged, so it does not change. As a result, the differential relation between liquidity and uncertainty as a function of transparency cannot be caused by transparency responding to liquidity.

To further assess the effect of volatility on the strength of the relation between transparency and liquidity, we disaggregate the *HVOL* indicator into quartiles. Results (not tabulated) provide evidence that the mitigating effect of transparency is monotonically increasing in the extremity of the uncertainty. These results are particularly interesting in light of the recent global economic crisis because they suggest that transparent firms are less likely to be affected by increases in investor uncertainty than are opaque firms.

3.4.3. Transparency and Firm-Level Governance. A final setting in which transparency is likely to be particularly important is where firm-level governance is weak, which can occur when ownership is highly concentrated and managers are thus more likely to be entrenched. To examine the effect of firm-level governance, we follow our earlier approach of including separate coefficients for low and high ownership concentration, but we now interact these piecewise coefficients with transparency. Results are reported in table 7. As before, we find that, at higher levels, increased ownership concentration appears to exacerbate agency problems and increase investor uncertainty. However, the interactions between ownership concentration and transparency remain significantly negative for both low and high ownership-concentration firms, suggesting that transparency is associated with less uncertainty irrespective of the ownership level. More important for our hypotheses, the effect of transparency is particularly strong among firms with the highest levels of ownership concentration, suggesting that, while transparency is associated with greater liquidity at all levels of ownership, the effects are particularly pronounced when ownership concentration (and, hence, the extent of expected agency problems) is high.

One final issue is whether firm-level governance interacts with country-level institutions. The idea is that concentrated ownership is likely to be particularly problematic when minority investor protection is weak and, hence, transparency is likely to be particularly important in those contexts. To assess this, we repeat the analysis, splitting between high- and low-*ASDI* firms, including the ownership concentration interaction. The last two columns of table 7 report the results of the analysis. Irrespective of country-level investor protection, transparency is more important for firms with more concentrated ownership. However, transparency matters most when concentrated ownership is combined with weak investor protection. The transparency coefficient is more than six times as large for observations with highly concentrated ownership in weak investor protection environments relative to observations with more diffuse ownership in strong investor protection environments.[35]

[35] In terms of economic significance, using a similar calculation procedure to that discussed in footnote 27, our results suggest that a shift in *TRANS* from the 25th to the 75th percentile is associated with a 32% decrease in zero return days for firms in the high %*CLHLD*/low *ASDI* subgroup and only a 5% decrease for firms in the low %*CLHLD*/high *ASDI* subgroup.

TABLE 7
Illiquidity and Firm-Level Governance

			ASDI			
	(1)		LOW		HIGH	
Variable	Estimate	*p*-value	Estimate	*p*-value	Estimate	*p*-value
Intercept	1.181	0.00	1.147	0.00	1.147	0.00
LNMVE	−0.088	0.00	−0.080	0.00	−0.093	0.00
BM	−0.003	0.00	−0.006	0.00	−0.001	0.34
LOSS	−0.011	0.00	−0.011	0.00	−0.013	0.00
STDRET	−0.274	0.00	−0.281	0.00	−0.271	0.00
ADREX	−0.036	0.00	−0.046	0.00	−0.032	0.01
ADR_NEX	−0.004	0.47	−0.048	0.00	0.015	0.02
%CLHLD_H	0.100	0.00	0.111	0.00	0.085	0.00
%CLHLD_L	−0.124	0.00	−0.135	0.00	−0.127	0.00
FUND_SMTHC	−0.012	0.01	−0.014	0.12	−0.015	0.01
%CLHLD_H TRANS*	−0.219***	0.00	−0.356**	0.00	−0.147***	0.00
%CLHLD_L TRANS*	−0.117	0.00	−0.243	0.00	−0.055	0.00
HIGH–LOW (*p*-value)			0.209 (0.00)			
HIGH–LOW (*p*-value)			0.188 (0.00)			
S.E. Clusters (#)			C-I(589)			
Fixed Effects	C,I,Y		C,I,Y		C,I,Y	
Observations	97,799		28,260		69,539	
S.E. Clusters (#)	F (15,681)		F (4,714)		F (10,967)	
Adj R^2	0.64		0.63		0.64	

The sample comprises firm-year observations between 1994 and 2007. The dependent variable in all models is *ILLIQ*, which is a combined measure of illiquidity equal to the average scaled percentile rank of the available liquidity measures, *ZERORET* and *LN(BIDASK)*. If *BIDASK* is unavailable, *ILLIQ* captures the percentile rank of *ZERORET*. *TRANS* is defined as the average of the scaled percentile rank of the variables: *ANALYST*, *R_ACCURACY*, *INTGAAP_S*, *BIG5*, and (1−*DIS_SMTHC*). If *R_ACCURACY* is unavailable, *TRANS* captures the average percentile rank of the remaining four measures. In columns 2 and 3, the sample is split between low *ASDI* and high *ASDI* depending on whether a firm's *ASDI* is above or below the sample median *ASDI*, where *ASDI* is the Anti-Self-Dealing Index developed by Djankov et al. [2008] for the country. The control variables are defined in table 2, panel A.

We include country (C), industry (I), and year (Y) fixed effects in all models, but do not report the coefficients. Industry fixed effects are based on the firm's two-digit ICB code. The table reports OLS coefficient estimates and *p*-values (two-sided) based on robust standard errors clustered at the firm-level. All continuous nonlogarithmic variables are truncated at the 1st and 99th percentiles. ***,(**) denotes that the difference between *%CLHLD* subgroups is significant at the 1% (5%) level. *HIGH–LOW* is equal to the coefficient difference on *%CLHLD* TRANS* across *ASDI* subgroups; *p*-values (two-sided) for these differences are presented in parentheses. Assessments of significance across partitions are made based on the standard errors (clustered at the country-industry level) of the coefficient on the interaction of *TRANS* and the partitioning variable in a fully interacted specification, i.e., we allow each variable, including the fixed effects, to vary by the partitioning variable.

$$\begin{aligned} ILLIQ_t = {} & \alpha_t + \beta_1 LNMVE_t + \beta_2 BM_t + \beta_3 LOSS_t + \beta_4 STDRET_t + \beta_5 ADR\_EX_t + \beta_6 ADR\_NEX_t \\ & + \beta_7 \%CLHLD\_H_t + \beta_8 \%CLHLD\_L_t + \beta_9 FUND\_SMTHC_t \\ & + \beta_{10} \%CLHLD\_H^* TRANS_t + \beta_{11} \%CLHLD\_L^* TRANS_t + FIXED\ EFFECTS + \varepsilon_t \end{aligned}$$

Overall, results from these analyses suggest important interactions between transparency and other aspects of the firm's environment. If overall uncertainty is high, either because of weak country-level institutions,

Repeating the analysis splitting based on country-level opacity using both our *DISCLOSE* and *MEDIA* variables, we find similar results (not tabulated for the sake of brevity).

economic volatility, weak firm-level governance, or a combination of these factors, we find that firm-level transparency appears to be particularly important to liquidity. Further, the fact that all of the interactions are consistent with expectations provides some comfort that our measures do, in fact, capture aspects of transparency and liquidity.

3.5 ENDOGENEITY AND REVERSE CAUSALITY

As noted earlier, a challenge for research on transparency is assessing causality. For example, the firm's choice of transparency may be a function of unobservable omitted variables (an endogeneity issue) or of the underlying liquidity in a firm's shares (a reverse causality issue). Our results are based on associations and, while we view them as descriptively interesting, it is inappropriate to draw strong conclusions about causality. That being said, we use several approaches to attempt to provide assurance that our results are not a product of endogeneity or reverse causality.[36]

First, our analyses include a wide range of controls which should capture many of the reasons that transparency might be endogenous. For example, suppose one is concerned that transparency and liquidity are both high because a firm trades on a U.S. exchange or because it attracts greater investor interest. Inclusion of variables such as cross listing, size, market-to-book, and growth should mitigate this concern. Similarly, our primary analyses use country, industry, and year fixed effects, which should control for country, industry, and time period factors.

Second, the robustness to inclusion of firm fixed effects makes it unlikely that our results simply reflect unmodeled cross-firm differences in transparency and liquidity. For example, if it were the case that some firms are of inherently higher liquidity and higher transparency, the relation between liquidity and transparency should not be evident in a within-firm comparison. If anything, our results are more consistent after controlling for cross-firm differences, suggesting that within-firm variation contributes meaningfully to the results.

Although the inclusion of firm fixed effects controls for time-invariant omitted variables, the possibility remains that firms experience a significant change not explicitly captured by our control variables, such as growth opportunities or financing needs, and that this change simultaneously affects both firm liquidity and our transparency proxies. To provide further assurance that such a change is unlikely to drive our results, as an untabulated robustness test we include several additional growth and financing variables in our analyses. The first is analysts' long-term growth forecasts from I/B/E/S. These forecasts, made during the firm's current fiscal year, capture analyst expectations of earnings growth over the next three to five years and, as such, should serve as a control for any predictable changes

[36] Tabulated results for additional analyses can be found in our Internet appendix, available at: https://sites.google.com/site/markgmaffett/research.

in the firm's growth opportunities. In addition, we include period $t + 1$ sales growth (thus implicitly assuming perfect foresight by the market). To control for shocks to a firm's financing needs, we employ two additional variables: capital raising and capital expenditures in period $t + 1$ (again assuming perfect foresight). Although these additional control variables substantially reduce our sample size and there is some slight attenuation of the coefficient on *TRANS*, our inferences throughout the paper remain unchanged by their inclusion, either individually or as a group.[37]

Third, the results are robust to the inclusion of lagged transparency. Untabulated results indicate similar conclusions when the contemporaneous transparency values are replaced by lagged values of transparency, again suggesting that our baseline results do not simply reflect an unobservable variable that drives both transparency and liquidity. Further, fixed effects results are robust to inclusion of lagged transparency, suggesting that innovations in transparency tend to be associated with higher subsequent liquidity.

Fourth, the analyses in the preceding section suggest that the effects are strongest following changes in investor uncertainty. Ideally, we would be able to observe exogenous shocks to transparency to help to identify the system. While it is difficult to identify exogenous transparency shocks, we can examine exogenous changes in country-level uncertainty over periods during which transparency is unlikely to change appreciably (e.g., one month windows). Given transparency is measured over the preceding year and is thus being held constant during those periods, and changes in liquidity remain a function of transparency, it is unlikely to be the case that the observed association between transparency and liquidity is simply the result of transparency changing in response to changes in liquidity.

Fifth, given that the association between liquidity and transparency varies predictably across subsamples, it is more difficult to envision a consistent theory in which causality is reversed yet the subsample results hold. For example, the theory would need to explain why, in low investor protection environments, in firms with higher ownership concentration, and in periods of high uncertainty, relatively lower levels of liquidity result in especially low levels of transparency.

Next, we attempt to assess endogeneity and causality by explicitly modeling transparency and liquidity in a two-stage least squares framework. Statistical approaches for dealing with simultaneity/causality concerns are not easy to implement because of the difficulty in obtaining appropriate instruments. In designing our system, we are guided primarily by analyst following research since we are not aware of attempts to model our other variables simultaneously. However, given that investor demand for information

[37] Because inclusion of analysts' long-term growth forecasts eliminates some countries from our sample entirely, and particularly those with weak country-level governance, we assess the robustness of our results to including this variable only in our primary specifications (tables 3 and 4).

provides the motivation for the choice of variables in modeling analyst following, such a motivation is likely to be shared when selecting related variables to model such as auditor choice, accounting standards, and earnings smoothing, and a similar estimation approach seems appropriate.[38]

In terms of analyst following, Roulstone [2003] provides some evidence that causality runs from analyst following to liquidity. He jointly estimates analyst following and liquidity for a sample of U.S. firms and concludes that analysts provide information to capital markets and do not simply chase liquidity. Similarly, Yu [2008] provides evidence that analysts discipline accounting choices around earnings management. In our two-stage least squares analysis (untabulated), we build on Roulstone [2003] and Yu [2008] and estimate a first-stage model that features transparency (*TRANS*) as a function of two sets of variables: potentially endogenous variables present in table 3 liquidity regression (*LNMVE, BM, STDRET, ADR_EX, ADR_NEX, %CLHLD_H, %CLHLD_L*, and *FUND_SMTHC*), and those suggested by research such as Lang and Lundholm [1996], Roulstone [2003], and Yu [2008] as instruments for transparency (return-earnings correlation and asset growth, both computed over the prior three- to five-year window, and one-year lagged return on assets).[39] Our second-stage model uses the same independent variables as the liquidity equations from column 5 of tables 3 and 4 and uses *ILLIQ* as the dependent variable. Analysis of the first-stage model suggests that our instruments are significantly related to transparency and the Kleibergen-Paap statistic indicates that we do not suffer from weak instruments (see Stock and Yogo [2005]). Results from the two-stage estimation are consistent with those reported earlier in that transparency remains significantly negatively correlated with illiquidity.[40]

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[38] It seems likely that, of our variables, the endogeneity of analyst following is of primary concern because, while it is intuitive that analysts might hesitate to follow firms with low liquidity given limited client interest, it is more difficult to envision how lower liquidity would motivate firms to choose lower quality auditors and accounting standards and to manage earnings, controlling for other factors (if anything, lower liquidity might provide incentives to improve reporting quality). Results are consistent if we treat only the *ANALYST* variable as endogenous.

[39] We selected these instruments because each is likely to be correlated with *TRANS* and we are unaware of theoretical reasons to expect them to be correlated with liquidity. While the resulting Hansen *J*-statistic supports this notion (i.e., we fail to reject the null hypothesis that our instruments are valid), such an overidentification test is only effective if at least one of the instruments is valid. Without a fully specified structural model, it is impossible to be sure these instruments satisfy the exclusion restriction and the 2SLS analysis should therefore be interpreted with caution.

[40] A related concern is self-selection bias because firms choose transparency levels. We implement a Heckman [1979] approach with a first-stage modeling transparency as a function of factors expected to both influence this decision and be exogenous to the second-stage illiquidity model (return-earnings correlation, asset growth and lagged return on assets as in our 2SLS analysis). The inverse Mills ratio is significant, indicating potential self-selection effects, but the coefficient on *TRANS* is not materially affected by its inclusion.

Subject to the caveat that dealing with endogeneity is difficult in these types of environments, the overall evidence presented in this section provides some confidence that our results do not simply reflect endogeneity or reverse causality.

3.6 OTHER ANALYSES

In this section, we discuss the results of several untabulated robustness tests also designed to bolster confidence in the interpretation of our results. First, because they represent such a significant portion of our sample, and thus threaten the generalizability of our results, we repeat our analyses eliminating Japanese and U.K. firms. Our inferences are robust to the exclusion of both of these countries. In fact, the relation between transparency and liquidity holds for 44 out of 46 sample countries, suggesting that the results are not driven by a small subset of countries. Similarly, to examine whether the results are sensitive to time period trends, we rerun the analysis by year. Results are consistent for each of the years in the sample.

Next, we implement two alternative estimation procedures for our discretionary smoothness variable, *DIS_SMTHC*. First, rather than pooling firm-year observations from all countries, we estimate *DIS_SMTHC* using the residuals from within-country regressions. We continue to find that discretionary smoothing is significantly negatively related to liquidity in all specifications. Second, we construct an alternative measure of *DIS_SMTHC* based on the predicted value from a regression of the residual value from the first-stage fundamental smoothness regression on determinants of discretionary smoothness discussed in appendix A. The residual from the fundamental smoothness regression consists of at least two components: discretionary smoothing and noise. This alternative estimation attempts to reduce noise in the discretionary component by capturing only the components of the residual related to likely determinants of discretionary smoothness. A limitation of this specification is that our first-stage fitted values are a linear function of a number of the other variables in the second-stage regression, so there is substantial multicollinearity. Including the fitted value in the liquidity regressions, the coefficient estimate on the fitted value remains strongly positive as predicted (p-value of 0.00) in all specifications; the fundamental smoothing variable remains significantly negative; and the analyst following, international GAAP, and accuracy measures remain significantly negative in all specifications. The Big-5 variable remains negative, but becomes insignificant in several specifications, likely reflecting multicollinearity.

Third, we include share turnover, defined as the annual volume of shares traded for the firm's fiscal year divided by the firm's total number of shares outstanding, as an additional control in the analysis. We do not include turnover in our primary analysis for two reasons. First, calculation of share turnover requires firm-level trading volume data, which materially reduces the size of our sample, and the missing observations are clustered in a

subset of country/years limiting the breadth of the sample. Second, it is not clear whether including a control for turnover is appropriate since turnover is itself a measure of liquidity (see, e.g., Lesmond [2005]) and reflects the effects of trading costs and infrequent trading. However, including share turnover as an additional control variable does not affect any of our conclusions.

Fourth, we consider regression specifications excluding *FUND_SMTHC* and *BM*. Our motivation for including these variables in our main analyses is to control for factors other than transparency, such as business risk, likely to affect liquidity. However, fundamental earnings smoothness and book-to-market have not typically been used as controls in liquidity regressions. Results excluding these variables are very similar to those presented in the tabulated analyses.

Fifth, we assess robustness to alternative constructions of the *TRANS* and *ILLIQ* variables. In our primary analyses, *TRANS* is based on the average percentile rank of *DIS_SMTHC*, *BIG5*, *ANALYST*, and *INTGAAP_S* and includes *R_ACCURACY* only where it is available. *ILLIQ* is similarly constructed as the average of the percentile rank of *ZERORET* and *BIDASK* when available. We construct the *TRANS* and *ILLIQ* variables in this fashion to preserve sample size and retain a broad cross section of firms and economic environments. To ensure that results are not driven by our treatment of missing observations, we re-estimate all of our analyses with modified *TRANS* and *ILLIQ* variables. The modified *TRANS* variable excludes *R_ACCURACY*, and is calculated as the average percentile rank of the four other proxies that are available for the entire sample. The modified *ILLIQ* variable is calculated only when both *ZERORET* and *BIDASK* are both nonmissing. Overall, results are very similar to our tabulated analyses, alleviating concerns that the construction of the *TRANS* variable unduly influences our results.

Finally, we include controls for the overall level and absolute value of accruals in our tests of liquidity effects. Jayaraman [2007] and Bhattacharya, Desai, and Venkataraman [2008] suggest that, in a U.S. setting, a higher level of accruals may be suggestive of greater informed trading and higher transaction costs. To ensure that our analysis is not affected by such circumstances, we replicate our analysis including the level and absolute value of accruals. We find that including either (or both) accruals measures has no effect on the inferences drawn in our liquidity analysis.

3.7 LINKING TRANSPARENCY AND LIQUIDITY TO VALUATION

We conclude our empirical section by specifically examining the relation between liquidity and both ex-ante cost of capital and firm valuation, as well as the role played by transparency. While it is true in theory that liquidity should affect cost of capital and valuation, as discussed earlier, there is little empirical evidence for an international set of firms on the significance or economic magnitude of this relation and, to our knowledge, there is no evidence that directly ties in the linkage with transparency.

We assess the effects of liquidity using two modeling approaches: implied cost of capital and Tobin's *Q*. There is debate in the empirical literature as to the optimal approach to estimating an implied cost of capital (see, e.g., Botosan and Plumlee [2005], Hail and Leuz [2006], and Lee, Ng, and Swaminathan [2009]). For this reason, we estimate four separate models frequently cited in the literature: 1) the modified PEG ratio model by Easton [2004]; 2) the Ohlson and Juettner-Nauroth [2005] model; 3) the Gebhardt, Lee, and Swaminathan [2001] model; and 4) the Claus and Thomas [2001] model. Following Hail and Leuz [2006], we take the average of these four models as our firm-specific measure of cost of capital.

We include the following control variables: *LNASSETS*, *LEV*, *STDRET*, *ADR_EX*, *ADR_NEX*, *BIAS*, and *RF_RATE* (the country-specific yearly risk-free interest rate). These controls are typical for the literature (see, e.g., Easton [2004], Botosan and Plumlee [2005], and Daske et al. [2008]). In addition, we include country, industry, and firm fixed effects.[41] Descriptive statistics for our cost of capital measure are reported in table 2, panel A. For our sample, the mean cost of capital is 11.4%, the median is 10.2%, and the interquartile range is from 8.0% to 13.3%. Overall, the relative order of magnitude seems reasonable based on prior studies and there is substantial variation among sample firms.

Our valuation regression estimates that feature implied cost of capital as the dependent variable are reported in table 8. It is important to note that the number of observations has dropped from 97,799 in the preceding analysis to 23,605 here, reflecting the limited availability of long-term analyst forecast data for our sample of firms. This is likely to bias against our finding strong results because the sample is limited to the largest, most heavily followed firms, which are likely to have fewer transparency issues and therefore a weaker relation between transparency and cost of capital. The table shows that cost of capital behaves as one would expect with respect to the control variables; cost of capital is higher the greater is the risk free rate, and is higher for smaller firms, more highly leveraged firms, and more volatile firms. Although not statistically significant at conventional levels, the exchange-listed ADR variable is negative, consistent with the prior literature.[42] More importantly, our variable of interest, *TRANS*, is negatively and significantly associated with cost of capital, which indicates that, as predicted, firms with higher levels of opacity tend to face a higher cost of equity capital. Results are consistent in the firm fixed effects estimations

[41] We exclude year fixed effects from the cost of capital analysis because, when included alongside country fixed effects, they are highly correlated with the risk-free rate measured by country each year. In additional analyses (untabulated), we confirm that results are robust to including country-year fixed effects in place of the risk-free rate.

[42] In an additional analysis (untabulated), we find that the coefficient on *ADR_EX* decreases by 25%–30% (i.e., it becomes more negatively associated with cost of capital), depending on the specification, when *TRANS* is excluded from the regression, suggesting that some of the benefits of cross listing may arise from increased transparency.

TABLE 8
Cost of Capital, Illiquidity, and Transparency

	(1)		(2)		(3)		(4)	
Variable	Estimate	*p*-value	Estimate	*p*-value	Estimate	*p*-value	Estimate	*p*-value
Intercept	0.172	0.00	0.149	0.00	0.164	0.00	0.150	0.00
LNASSETS	−0.003	0.00	−0.002	0.00	−0.009	0.00	−0.007	0.00
LEV	0.019	0.00	0.018	0.00	0.018	0.00	0.017	0.03
STDRET	0.084	0.00	0.087	0.00	0.026	0.00	0.026	0.00
ADR_EX	−0.003	0.13	−0.003	0.19	−0.004	0.29	−0.004	0.29
ADR_NEX	0.000	0.93	0.000	0.89	0.000	0.95	0.000	0.93
BIAS	0.089	0.00	0.086	0.00	0.059	0.00	0.059	0.01
RF_RATE	0.127	0.00	0.114	0.00	0.162	0.00	0.158	0.02
ILLIQ			0.021	0.00			0.012	0.00
TRANS	−0.041	0.00	−0.032	0.00	−0.017	0.00	−0.013	0.00
Mediating Effect (*p*-value)			0.009 (0.00)				0.004 (0.00)	
Fixed Effects	C, I		C, I		F		F	
Observations	23,605		23,605		23,605		23,605	
S.E. Clusters (#)	F (5,387)		F (5,387)		F (3,387)		F (5,387)	
Adj. R^2	0.28		0.28		0.52		0.52	

The sample comprises firm-year observations between 1994 and 2007. The dependent variable is a firm's average implied cost of capital, *AVG_COC*, calculated as the yearly average of four commonly used empirical estimation techniques as described in table 2, panel A. *TRANS* is defined as the average of the scaled percentile rank of the variables: *ANALYST*, *R_ACCURACY*, *INTGAAP_S*, *BIG5*, and (1−*DIS_SMTHC*). If *R_ACCURACY* is unavailable, *TRANS* captures the average percentile rank of the remaining four measures. *ILLIQ* is a combined measure of illiquidity equal to the average scaled percentile rank of the available liquidity measures, *ZERORET* and *LN(BIDASK)*. If *BIDASK* is unavailable, *ILLIQ* captures the percentile rank of *ZERORET*. *Mediating Effect* is equal to the decrease in the coefficient on *TRANS* that arises from including *ILLIQ* in the cost of capital model. *P*-values (two-sided) for the significance of the mediating effect, based on a Sobel [1982] test, are presented in parentheses. The control variables are defined in table 2, panel A.

We include country (C) and industry (I) fixed effects in models 1 and 2, and firm (F) fixed effects in models 3 and 4, but do not report the coefficients. Industry fixed effects are based on the firm's two-digit ICB code. The table reports OLS coefficient estimates and *p*-values (two-sided) based on robust standard errors clustered at the firm-level. All continuous nonlogarithmic variables are truncated at the 1st and 99th percentiles.

$$AVG\_COC_t = \alpha_t + \beta_1 LNASSETS_t + \beta_2 LEV_t + \beta_3 STDRET_t + \beta_4 ADR\_EX_t + \beta_5 ADR\_NEX_t + \beta_6 BIAS_t + \beta_7 RF\_RATE_t + \beta_8 ILLIQ_t + \beta_9 TRANS + FIXED\ EFFECTS + \varepsilon_t$$

(column 3), suggesting that the analysis does not simply capture differences in underlying characteristics across firms.

Beyond documenting that transparency is negatively associated with cost of equity capital, we also examine whether liquidity is an important channel for this effect using a mediation analysis (e.g., Hammersley [2006]).[43] The results of the mediation analysis (with country, industry, and year fixed

[43] The goal of a mediation analysis is to assess whether one variable (e.g., liquidity) serves as a significant channel through which another variable (e.g., transparency) affects a particular dependent variable (e.g., cost of equity capital). The first step is to show that the dependent variable of interest (e.g., *AVG_COC*) is related to the dependent variable (*TRANS*). Next, the expected mediator variable (*ILLIQ*) is shown to be related to the independent variable (*TRANS*). Third, the mediator variable (*ILLIQ*) is shown to be correlated with the dependent variable (e.g., *AVG_COC*). Finally, the mediator (*ILLIQ*) and the original independent variable (*TRANS*) are included in the same regression along with the dependent variable of interest (e.g., *AVG_COC*). If the mediator variable mediates the relation between the dependent

effects in column 2 and with firm fixed effects in column 4) confirm our predictions. Specifically, the coefficient on *TRANS* decreases significantly when *ILLIQ* is added to the regression analyses, suggesting that liquidity is a significant channel through which transparency affects cost of capital.[44] Moreover, we can use the coefficients on our variables to broadly gauge economic significance. They indicate that a shift from the 75th to the 50th illiquidity percentile is associated with a decrease in cost of capital of just less than half of a percentage point (45 basis points) and that a shift from the 75th to the 50th transparency percentile is associated with an increase in cost of capital of approximately one-third of a percentage point (32 basis points). These results suggest that the liquidity and transparency effects we document are economically important as well as statistically significant.[45]

We also estimate the relation between liquidity and valuation as measured with Tobin's *Q*. While a cost of capital approach is more direct, it limits our analysis to a small subsample of the largest firms and relies crucially on the assumptions around analyst forecasts. This is particularly an issue in international settings where analyst following, when it exists, tends to be limited. For our Tobin's *Q* analysis, the sample size increases substantially to 76,936 and includes a much wider range of firms.[46]

Tobin's *Q* is defined as: (book value of assets + (market value of equity – book value of equity))/book value of assets. It is designed to reflect the valuation placed on the assets by the market relative to their book value and inherently incorporates the cost of capital used in discounting future cash flows. Table 2, panel A, provides statistics for Tobin's *Q* (*Q*). The median *Q* is about 1.10, indicating that investors value assets slightly above their book value.

Table 9 reports results for Tobin's *Q*. We include control variables from the prior literature, along with country, industry, year and firm fixed effects.[47] Results for the controls are consistent with prior literature. Tobin's *Q* tends to be higher for firms that are smaller, more profitable, more highly levered, have higher growth, are cross listed on U.S. exchanges, and pay dividends. In terms of our variable of interest, the coefficient on

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variable and the original independent variable, then the significance of the original independent variable will be reduced over the first-stage regression and the mediator variable will be significant.

[44] The significance of a mediation effect is assessed with a Sobel test following Sobel [1982].

[45] Economic significance for cost of capital is calculated as follows: from table 2, panel A, *COC_AVG* has a median value of 0.102 and an interquartile shift in *ILLIQ* leads to an increase of 0.220 (=0.720−0.500). The coefficient on *ILLIQ* in table 8, column 2 is 0.021. Economic significance is calculated as ((0.102 − (0.220∗0.021))/0.102) −1. Economic significance for the transparency effect is calculated using an identical procedure.

[46] The sample size in our Tobin's *Q* analysis is smaller than in our main analysis because we require data on capital expenditures, which are missing for approximately 20,000 firms. All results are robust to assuming that missing capital expenditures are equal to zero and estimating the models using the full sample.

[47] See, for example, Claessens et al. [2002], Lins [2003], and Kalcheva and Lins [2007].

TABLE 9
Firm Value, Illiquidity and Transparency

	(1)		(2)		(3)		(4)	
Variable	Estimate	*p*-value	Estimate	*p*-value	Estimate	*p*-value	Estimate	*p*-value
Intercept	0.641	0.00	1.646	0.00	1.825	0.00	2.324	0.00
LNASSETS	−0.053	0.00	−0.124	0.00	−0.200	0.00	−0.235	0.00
LEV	0.410	0.00	0.493	0.00	0.544	0.00	0.610	0.00
CASH_TA	1.018	0.00	0.958	0.00	0.527	0.00	0.535	0.00
SG	0.282	0.00	0.252	0.00	0.215	0.00	0.202	0.00
NIEX_TA	1.738	0.00	1.668	0.00	1.335	0.00	1.300	0.00
DIV_DUM	−0.038	0.00	−0.050	0.00	0.050	0.00	0.034	0.00
CAPEX_TA	1.340	0.00	1.238	0.00	1.068	0.00	0.998	0.00
ADR_EX	0.150	0.00	0.098	0.00	0.061	0.13	0.033	0.42
ADR_NEX	0.108	0.00	0.095	0.00	0.010	0.80	0.006	0.88
ILLIQ			−0.811	0.00			−0.515	0.00
TRANS	0.940	0.00	0.707	0.00	0.358	0.00	0.288	0.00
Mediating Effect (*p*-value)			0.233 (0.00)				0.07 (0.00)	
Fixed Effects	C,I,Y		C,I,Y		F,Y		F,Y	
Observations	76,936		76,936		76,936		76,936	
S.E. Clusters (#)	F (13,466)		F (13,466)		F (13,466)		F (13,466)	
Adj. R^2	0.24		0.29		0.68		0.69	

The sample comprises firm-year observations between 1994 and 2007. The dependent variable is Tobin's *Q*, calculated as total assets less book value plus market value scaled by total assets. *TRANS* is defined as the average of the scaled percentile rank of the variables: *ANALYST*, *R_ACCURACY*, *INTGAAP_S*, *BIG5*, and (1−*DIS_SMTHC*). If *R_ACCURACY* is unavailable, *TRANS* captures the average percentile rank of the remaining four measures. *ILLIQ* is a combined measure of illiquidity equal to the average scaled percentile rank of the available liquidity measures, *ZERORET* and *LN(BIDASK)*. If *BIDASK* is unavailable, *ILLIQ* captures the percentile rank of *ZERORET*. *Mediating Effect* is equal to the decrease in the coefficient on *TRANS* that arises from including *ILLIQ* in the *Q* model. *P*-values (two-sided) for the significance of the mediating effect, based on a Sobel [1982] test, are presented in parentheses. The control variables are defined in table 2, panel A.

We include country (C), industry (I), and year (Y) fixed effects in models 1 and 2, and firm (F) and year (Y) fixed effects in models 3 and 4, but do not report the coefficients. Industry fixed effects are based on the firm's two-digit ICB code. The table reports OLS coefficient estimates and *p*-values (two-sided) based on robust standard errors clustered at the firm-level. All continuous nonlogarithmic variables are truncated at the 1st and 99th percentiles.

$$Q_t = \alpha_t + \beta_1 LNASSETS_t + \beta_2 LEV_t + \beta_3 CASH\_TA_t + \beta_4 SG_t + \beta_5 NIEX\_TA_t + \beta_6 DIV\_DUM_t + \beta_7 CAPEX\_TA_t + \beta_8 ADR\_EX_t + \beta_9 ADR\_NEX_t + \beta_{10} ILLIQ_t + \beta_{11} TRANS + FIXED\ EFFECTS + \varepsilon_t$$

TRANS is strongly positive, suggesting that investors place higher valuations on more transparent firms.

Again, we estimate a mediation analysis to disentangle the valuation effect of transparency that occurs through liquidity from the direct effect of transparency on valuation. The mediation analysis suggests that transparency is important to valuation, both through the liquidity channel as well as through other channels. Including *ILLIQ* in the regression significantly mediates the effect of transparency on valuation, although the effect is only partial because the coefficient on *TRANS* remains significantly positive. Finally, the relation between liquidity and *Q* appears to be economically meaningful as well, with a shift from the 75th to the 50th percentile of

ILLIQ associated with about a 16% increase in Tobin's *Q* and a similar shift in *TRANS* associated with a 6.5% decrease in Tobin's *Q*.[48]

4. Conclusions

Reductions in the liquidity and valuation of securities traded in global capital markets during the recent financial crisis have demonstrated the importance of understanding more fully the drivers of a firm's stock market liquidity and associated linkages to valuation. In this paper, we examine whether reduced transparency is associated with increased transaction costs and lower liquidity in a firm's shares and, therefore, increased cost of capital and reduced valuation. We also investigate the extent to which the relation between transparency and liquidity is influenced by institutional and firm-level factors and by time series variation in uncertainty.

For a global sample of firms, our evidence suggests that increased transparency, as reflected in reduced earnings management, higher quality auditing, a serious commitment to international accounting standards, increased analyst following, and smaller analyst forecast errors, is associated with lower bid-ask spreads and greater liquidity. The relation is particularly pronounced in environments in which there is likely to be more inherent uncertainty (weak country-level institutions, time periods of increased country-level volatility, and when ownership is concentrated), suggesting that firm-level transparency is most important in the presence of other informational issues. Our results also provide evidence that liquidity represents an important channel through which transparency becomes associated with a lower cost of capital and higher valuation. Taken together, our results suggest that a focus on the transparency provided to those who invest in a firm's securities could be a fruitful component of an effort to more fully understand the increases in illiquidity and decreases in valuation for many assets worldwide associated with the recent financial crisis.

Our results are subject to several caveats. First, we focus on only one potential consequence of increased transparency, improved liquidity. Of course, increased transparency entails other costs and benefits. As a consequence, our results do not imply that managers would be better off by increasing transparency, only that benefits may accrue through reduced transaction costs and increased liquidity. It is possible that other costs associated with increased transparency more than offset the liquidity benefits. There is room for future research examining more specifically the tradeoffs in establishing an optimal transparency level.

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[48] Economic significance for Tobin's *Q* is calculated as follows: from table 2, panel A, *Q* has a median value of 1.101 and an interquartile shift in *ILLIQ* leads to an increase of 0.220 (=0.720–0.500). The coefficient on *ILLIQ* in table 9, column 2 is equal to −0.811 and economic significance is calculated as ((1.101 − (0.220*−0.811))/1.101) −1. Economic significance for the transparency effect is calculated using an identical procedure.

Second, as discussed earlier, it is difficult to ascertain causality. Our analyses are based on associations, and we cannot be certain to what extent the relations are causal. While we attempt to control for a wide range of potentially important factors and to account for possible endogeneity, conclusions should be drawn with caution. There is substantial scope for additional research identifying more specifically the channels through which transparency may affect liquidity. Overall, though, we view our paper as providing interesting initial evidence on the potentially important effects of transparency on liquidity and valuation in a global setting.

APPENDIX A
Computation and Tests of the Discretionary Earnings Management Proxy: DIS_SMTHC

We compute *DIS_SMTHC* based on two earnings smoothness measures commonly used in the literature. The first earnings smoothness measure (*SMTH1*) captures the volatility of earnings relative to the volatility of cash flows with the idea being that, the more firms use accruals to manage earnings, the smoother net income will be relative to cash flows (Leuz, Nanda, and Wysocki [2003] and Francis et al. [2004]). *SMTH1* is measured as the standard deviation of net income before extraordinary items divided by the standard deviation of cash flow from operations, where net income before extraordinary items and cash flow from operations are scaled by average total assets and the standard deviations are calculated using rolling time intervals requiring a minimum of three and a maximum of five years of data. Cash flow from operations is equal to net income before extraordinary items minus accruals, where accruals are defined as the change in current assets minus the change in current liabilities minus the change in cash plus the change in current debt in current liabilities minus depreciation and amortization expense.

The second earnings smoothness measure (*SMTH2*) is the correlation between the cash flow from operations scaled by total assets and total accruals scaled by total assets. The idea behind this measure is that, to the extent managers create accrual reserves in good times and use them to compensate for poor cash flows in bad times, accruals and cash flows will be more negatively correlated (Lang, Raedy, and Wilson [2006], Barth, Landsman, and Lang [2008]). We note that Leuz et al. [2003] and Bhattacharya, Daouk, and Welker [2003] calculate their correlation-based measure using the change in cash flows from operations and the change in total accruals, whereas our correlation measure is based on the level. We draw identical inferences when defining *SMTH2* based on changes; however, the sample sizes are smaller due to the additional data requirements of the change measures. Both smoothing measures (*SMTH1*, *SMTH2*) are multiplied by negative one so that larger values represent firms with smoother earnings.

The smoothness of earnings relative to cash flows is clearly a natural function of the fundamentals that affect a firm's operating environment, but we are interested in the portion in excess of naturally occurring earnings smoothness. As a consequence, we draw from prior research on the determinants of earnings smoothness and specify an equation designed to capture, to the extent possible, the expected level of earnings smoothness for a firm. We then measure discretionary (excess) smoothing using the residual from the regression specified below:

$$\begin{aligned} SMTH_t = {} & \beta_1 LNASSETS_t + \beta_2 LEV_t + \beta_3 BM_t + \beta_4 STD\_SALES_t \\ & + \beta_5 \%LOSS_t + \beta_6 OPCYCLE_t + \beta_7 SG_t + \beta_8 OPLEV_t \\ & + \beta_9 AVECFO_t + \sum_{a=1}^{18} \alpha_a IND_i + \sum_{b=1}^{14} \alpha_b YEAR_i + \varepsilon_t \end{aligned}$$

The right-hand side variables are: *LNASSETS*, the log of total assets measured in millions of U.S. dollars, a measure of firm size; *LEV*, total debt divided by total assets, to capture differences in financing choices; *BM*, the ratio of book value to market value of equity, to reflect the extent of the firm's intangible assets and expected earnings growth; *STD_SALES*, the standard deviation of sales, to capture the volatility of a firm's underlying operating environment; *%LOSS*, the proportion of years that a firm experiences losses over the last three to five years, to capture differences in the accruals properties of loss observations; *OPCYCLE*, the log of days of accounts receivable plus inventories, to capture the length of the firm's operating cycle; *SG*, the average sales growth over the past three to five years, to capture growth opportunities; *OPLEV*, net property, plant and equipment divided by total assets, to capture capital intensity; *AVECFO*, average cash flow from operations divided by total assets measured over the last five years, to capture a firm's general level of profitability; and indicator variables for a firm's industry because the properties of accruals are likely to depend on industry, as well as year indicator variables to control for macro economic cycles that could affect earnings cycles.

After we obtain each of the two discretionary smoothness regression residual measures for *SMTH1* and *SMTH2*, they are then scaled into percentile ranks, and combined by taking the average. This variable is referred to as *DIS_SMTHC* and is used to proxy for earnings management in our liquidity regressions. We follow the same procedure in computing our proxy for fundamental earnings smoothness, which is used as a control variable in our tests. The predicted values obtained from the model for *SMTH1* and *SMTH2* are scaled into ranks and averaged—this variable is called *FUND_SMTHC*.

Because earnings management is inherently difficult to measure, we next conduct a set of tests designed to build confidence in the selection of our residuals-based measure of earnings management. For these tests, we benchmark our *SMTH1* and *SMTH2* measures against likely determinants

of discretionary earnings management. If our measures reflect managerial discretion, there should be predictable correlations with managerial incentives to smooth earnings and with institutional constraints on those incentives that go beyond the correlations with fundamentals-based variables that we have specified in our model above.

Therefore, we conduct an additional analysis that assesses the association between our smoothing measures and a set of incentive and oversight proxies. As country-level measures of governance, we include either the recently developed Anti-Self-Dealing Index (*ASDI*) of Djankov et al. [2008], which has been shown to be particularly useful in determining the extent to which managerial self-dealing is likely to be controlled by a country's institutional factors or *DISCLOSE* from La Porta, Lopez-de-Silanes, and Shleifer [2006], which summarizes the disclosure requirements faced by firms in a given country. In addition, we include an indicator variable for the degree of alignment between tax and financial reporting (*TXBKCONFORM* from Ashbaugh and LaFond [2004]) since, in countries with a high degree of alignment, the incentives managers face to smooth earnings for taxes will carry over to smoother accounting earnings (Alford et al. [1993], Ali and Hwang [2000], and Kasanen, Kinnunen, and Kiskanen [1996]).

In terms of firm-level determinants, we include an indicator variable for whether the firm is listed on a U.S. exchange, *ADR_EX*, since the U.S. regulatory environment is considered one of the most demanding in the world. We note that, while firms trading in U.S. markets are not required to report local accounts that comply with U.S. GAAP, Pownall and Schipper [1999], Ashbaugh and Olsson [2002], and Lang, Ready, and Wilson [2006] suggest that non-U.S. firms required to prepare U.S. GAAP financial information choose alternatives under IFRS or their domestic standards that are closer U.S. GAAP. We also include an indicator, *ADR_NEX,* for other types of U.S. cross listings (Level 1 and Rule 144A listings) that allow access to U.S. investors, but do not commit the firm to SEC registration requirements since firms may choose to curtail discretionary smoothing to enhance the informativeness of their accounting earnings by U.S. investors, even if they are not subject to additional regulatory oversight. Oversight by informational intermediaries likely affects firms' incentives and ability to smooth earnings as well. We use analyst following (*ANALYST*) as a proxy for the demand for transparent financial information by capital market participants.[49] Further, because larger auditing firms are likely to have greater resources and greater legal and reputational exposure, we expect attestation by a Big-5 auditing firm (*BIG5*) to be associated with less discretionary smoothing (Fan and Wong [2005]). Finally, following Bradshaw and Miller [2008] and Barth, Landsman, and Lang [2008], we expect better accounting standards to reduce the ability to manage earnings and include an indicator, *INTGAAP,* for firms that have adopted either IAS or U.S. GAAP.

[49] All of our results are consistent if we remove analyst following from the model, given the potential simultaneity discussed in the body of the paper.

Results for our earnings smoothing measures (not tabulated) are all consistent with predictions. In particular, our measures of earnings management are lower for firms in countries with better investor protection and a weaker link between tax and financial reporting, and in firms with higher analyst following and a Big-5 auditor that report under IFRS or U.S. GAAP in their local accounts and trade in the U.S., particularly if they trade on a U.S. exchange. Taken together, these results provide some comfort that our smoothing measures behave as though they reflect managerial discretion in the sense that they are positively correlated with incentives to manage earnings and negatively correlated with impediments to earnings management.

APPENDIX B
Results Disaggregating the Transparency Variable

Throughout the analyses of the interactive transparency effects (i.e., tables 5–7), for parsimony we report all results using an aggregate measure of transparency. This treatment corresponds to the notion that our variables likely capture the same latent firm characteristic, transparency, and that they are unlikely to be independent of each other.[50] Although our primary interest is overall transparency, as opposed to the individual components, in this appendix we report results disaggregating transparency into its components (i.e., *DIS_SMTHC*, *BIG5*, *ANALYST*, *R_ACCURACY*, and *INTGAAP_S*).[51] In general, conclusions are consistent across the components of the transparency variable, although the relations for the subcomponents are not always statistically significant.

Beginning with panel A of table 5, for the *ASDI* split, each of the individual transparency variables is more negatively related to illiquidity in the low *ASDI* group, except for *BIG5*, where the difference between the subgroups is insignificant. For the *DISCLOSE* split, *ANALYST*, *R_ACCURACY* and *INTGAAP_S* are more negatively related to *ILLIQ* in the low *DISCLOSE* group, while *DIS_SMTHC* and *BIG5* are insignificantly different between the groups. For the *MEDIA* split, the difference between the low and high subgroups is significant for the *ANALYST, DIS_SMTHC*, and *INTGAAP_S* variables but not for *BIG5* and *R_ACCURACY*. In panel B of table 5, disaggregation of the transparency variable shows that each of the individual variables is most important to illiquidity in the country grouping with the worst governance ($GOV = 0$) and least important in the group with the

[50] For example, accounting standards choice likely affects ability to manage earnings, analyst following, forecast accuracy and even choice of auditor, so it is difficult to separate out the individual component effects.

[51] In all of the analyses in this appendix, assessments of statistical significance are based on an identical specification to that reported for the corresponding aggregate specification in the primary analyses in terms of both the type of fixed effects included and the dimension(s) along which standard errors are clustered.

best governance ($GOV = 3$), except for the *DIS_SMTHC* and *BIG5* variables, where these differences are not statistically significant.

In table 6, across all three columns, when disaggregating the *TRANS* variable into its individual components, we find that *DIS_SMTHC*, *BIG5*, and *ANALYST* are all significantly more negatively related to *ILLIQ* following periods of high uncertainty. While the interactions between *HVOL* and *R_ACCURACY* and *INTGAAP_S* generally maintain their negative signs, the coefficients are insignificant.

In terms of the closely held shares interactions in table 7, column 1, *DIS_SMTHC* and *ANALYST* are both most negatively related to *ILLIQ* when there is a larger proportion of shares that are closely held, while the remaining three variables are insignificantly different between the partitions. Splitting the sample into high and low *ASDI* groups, all of the variables, except for *DIS_SMTHC*, are most negatively related to illiquidity when more shares are closely held and protection against self-dealing is weak.

Following a similar logic to that used in constructing our aggregate transparency variable, we also use an aggregate illiquidity variable in tables 5–7, combining the illiquidity proxies *ZERORET* and *BIDASK*. All results in these analyses are robust to using the *ZERORET* measure. The same is true for *BIDASK* except in the following cases: (1) In table 5, panel B, the difference in the magnitude of the *TRANS* coefficient is larger in the low *DISCLOSE* subsample, but is insignificantly different from the high *DISCLOSE* coefficient, and (2) in table 6, column 3, the interaction between *TRANS* and *HVOL* is negative, but not significant.

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