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Liquidity, Liquidity Risk and Stock Returns: Evidence from Japan

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Abstract

This paper investigates whether liquidity and liquidity risk are priced in Japan. Using modified Amihud illiquidity measures, we find both cross-sectional and time series evidence that liquidity is priced in the Japanese stock market during the period 1975–2006. The evidence is largely consistent with Amihud's (2002) findings in the US market. We further employ the liquidity-adjusted CAPM proposed by Acharya and Pedersen (2005) to examine whether liquidity risk is priced in Japan. Consistent with Acharya and Pedersen's findings in the US, we show that liquidity risk is priced in the stock market, in addition to the liquidity level. These findings strengthen the confidence that liquidity is a determinant of stock returns.

Keywords: *liquidity*, *Amihud illiquidity measure*, *liquidity-adjusted CAPM*, *Japan*

JEL classification: *G*12

1. Introduction

Liquidity (or illiquidity) and liquidity risk have important practical as well as academic implications. Several studies have documented that liquidity can explain the difference in

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the cross section of stock returns in the USA.[1] Using the illiquidity measure proposed by Amihud (2002) and the liquidity-adjusted capital asset pricing model (CAPM) proposed by Acharya and Pedersen (2005), we examine whether liquidity and liquidity risk are priced in the Japanese stock market.

Over the sample period analysed in this paper, Japan had the second-largest stock market in the world, next to the USA, in terms of both market capitalisation and number of listed securities.[2] As pointed out by Chan *et al.* (1991), the study of the Japanese stock market alongside that of the USA is important for the evaluation of empirical models of cross-sectional stock returns. Empirical confirmation of the same determinants in both countries would strengthen confidence in the model, while differences between countries would induce further exploration of asset pricing theories. In addition, evidence from the Japanese market may shed light on the subsumption of explanatory variables and robustness with regard to time period and sample selection.

Chan *et al.* (1991) relate cross-sectional difference in stock returns in Japan to earnings yield, firm size, book-to-market ratio (BM), and cash flow yield. The authors uncover significant relations between these variables and expected returns in the Japanese market, which is consistent with US findings. Daniel *et al.* (2001) use Japanese data to test the robustness of Daniel and Titman's (1997) empirical asset pricing model based on US data. Employing market microstructure data from Japan, Lehmann and Modest (1994) analyse liquidity on the Tokyo Stock Exchange (TSE) and compare their findings with those obtained on the New York Stock Exchange (NYSE). Bremer and Hiraki (1999) find evidence linking short-term returns of individual TSE stocks to lagged trading volume and compare their results with those found in the US stock market. Using market microstructure data, Ahn *et al.* (2002) find that stock returns on the TSE are positively related to illiquidity measures.

Few studies have examined the relation between stock returns and liquidity in Japan. An exception is the paper by Chang *et al.* (2010). In a cross-sectional setup, these authors examine the relations between stock returns and six monthly liquidity measures, including such traditional measures as turnover and dollar trading volume, Amihud's (2002) illiquidity measure, the proportion of trading days with zero returns proposed by Lesmond *et al.* (1999), and the liquidity measure proposed by Liu (2006). Chang *et al.* also analyse the relation between stock returns and liquidity variability (a rough proxy for liquidity risk) measured by the coefficients of variation for turnover, the dollar trading volume, and the Amihud illiquidity measure. Chang *et al.* (2010) obtain mixed results: They show that stock returns are negatively related to four out of six liquidity measures for the period 1975–2004, but not consistently so across the subsample periods. They also find that liquidity variability measures are related to stock returns, but with the wrong sign; that is, the more variable the liquidity, the lower the stock return.

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[1] See Amihud and Mendelson (1986), Glosten and Harris (1988), Hasbrouck (1991), Brennan and Subrahmanyam (1996), Eleswarapu (1997), Chalmers and Kadlec (1998), Easley *et al.* (2002), and Pastor and Stambaugh (2003), among others. Using mainly market microstructure data from the USA and various estimation techniques, these authors report a positive relation between illiquidity and stock returns across companies. Subrahmanyam (2010) provides a whole section to review the literature about liquidity as a cross-sectional predictor of stock returns. Neumann (2003) documents that liquidity has an impact on price spreads between dual-class shares.

[2] China overtook Japan as the world's second largest stock market in terms of market capitalisation in the middle of 2009 .

Our study is different from that of Chang *et al.* (2010) in the following aspects. First, we focus on the Amihud illiquidity measure. Various proxies for liquidity have been used in previous studies and, as pointed out by Lesmond (2002), some traditional proxies, such as turnover, trading volume, and firm size, may capture the effects of variables not related to liquidity. On the other hand, some finer and more accurate measures based on market microstructure data, such as bid–ask spread, amortised effective bid–ask spread, price response to signed order flow, and the probability of information-based trading, are not generally available, especially over long periods of time. Amihud (2002) develops a measure that aims to balance the limits of data availability and accuracy: the average of the daily ratio of the absolute stock return to its dollar trading volume over a given period. This measure is simple, since it only requires the input of daily data to construct and is applicable to all securities and time periods.[3] It is also intuitive, since it can be interpreted as the daily stock price response to $1 of trading volume, which is consistent with Kyle's (1985) concept of illiquidity, that is, the response of price to order flow, and Silber's (1975) thinness measure, that is, the ratio of the absolute price change to the absolute excess demand for trading. After comparing a few alternative liquidity measures constructed using daily data, Hasbrouck (2005) concludes that the Amihud measure works the best. This study further modifies the Amihud measure based on the suggestions of Hasbrouck (2005) and Acharya and Pederson (2005), which should make the Amihud measure better reflect illiquidity in cross-sectional and time series analyses.

Second, we examine the time series relation between liquidity and stock returns, which is not examined by Chang *et al.* (2010). In fact, not many studies examine the relation between liquidity and stock returns over time.[4] As pointed out by Amihud (2002), this is probably due to the fact that illiquidity measures based on microstructure data for long time periods are not available in most markets. Using Amihud's (2002) approach, we test whether stock returns in Japan are related to the expected and unexpected illiquidity over time. In addition, we examine whether there is the flight to liquidity effect.

Third, we capitalise on Acharya and Pedersen's (2005) work to investigate whether liquidity risk is related to stock returns. These authors argue that not only the level of liquidity but also the uncertainty of liquidity (liquidity risk) should be priced. However, using liquidity variability measures, Chordia *et al.* (2001) and Chang *et al.* (2010) find a negative and strong relation between liquidity uncertainty and expected stock returns, which is counterintuitive. In contrast, Acharya and Pedersen (2005) propose a liquidity-adjusted CAPM. Using the Amihud illiquidity measure and stock returns to compute liquidity betas, Acharya and Pedersen find some evidence that illiquidity betas are priced in the USA and that a liquidity-adjusted CAPM is better than the standard CAPM in terms of goodness of fit. Our study using Japanese data provides an out-of-sample test of the authors' liquidity-adjusted CAPM.

We also take into consideration factors unique to the Japanese market in designing our test and examining alternative model specifications. Using monthly data from 1975 to 2006, our empirical findings are largely consistent with those found in Amihud (2002) and Acharya and Pedersen (2005). First, the illiquidity level is positively related to

[3] Pastor and Stambaugh (2003) use daily data to construct a liquidity measure based on signed order flow. Lesmond (2002) develops a liquidity estimate based on the percentage of zero return trading days over a certain period, such as a year.

[4] Bekaert *et al.* (2003) use pooled data to study the relation between liquidity and expected returns across firms over time in emerging markets.

stock returns across firms for the whole sample period. Second, the expected liquidity is positively related to stock returns while the unexpected liquidity is negatively related to stock returns over time. We also find some evidence in support of the 'flight to liquidity' hypothesis proposed by Amihud (2002). Finally, liquidity risk is priced, in the sense that the liquidity-adjusted CAPM fits data better than the standard CAPM. However, a puzzling finding is that liquidity seems unrelated to stock returns across firms for the sub-sample period 1990–2006. A similar finding is also documented by Chang *et al.* (2010).

The rest of the paper is organised as follows. Section 2 describes the data and the illiquidity measure. Section 3 examines the cross-sectional relation between illiquidity and stock returns, while the time series relation between illiquidity and stock returns is studied in Section 4. Section 5 further examines whether illiquidity risk is priced in the liquidity-adjusted CAPM setting, and Section 6 concludes.

2. Data and the Illiquidity Measure

All the data used in this study, unless otherwise stated, are from the Pacific-Basin Capital Markets Database (PACAP) Japan Database, which covers the period from January 1975 to December 2006.[5] According to Amihud (2002), daily stock illiquidity is computed as the ratio of absolute daily returns to the daily trading value. However, we follow Hasbrouck (2005) by further taking the square root of this ratio to filter out the impact of outliers:

$$ILL_d^i = \sqrt{|R_d^i|/VAL_d^i} \tag{1}$$

where R_d^i is the return for stock i on day d, VAL_d^i is the trading value for stock i on day d in millions of yen, and ILL_d^i represents the absolute percentage price change per ¥1 million of trading value with the square root transformation.

Following Lehmann and Modest (1994) and Bremer and Hiraki (1999), we include only stocks listed on the first section of the TSE, because the stocks in different sections satisfy different listing criteria and are likely to have very different trading and liquidity characteristics. For example, stocks listed on the first section are much larger and much more actively traded than those in the second section. We use daily stock returns and trading values in each month to construct our monthly illiquidity series. Our sample period starts July 1975 and ends in December 2006 (378 monthly observations). It starts in July 1975 instead of January 1975 for the following three reasons: First, most Japanese firms use March as their fiscal year-end, and financial reports may not be available until June, and we need some of the annual report data to construct proxies for firm characteristics, such as the BM and cash flow-to-price ratio (CP). Second, we use previous 150 daily return data to compute the monthly beta. Third, we use the average monthly returns for the previous two quarters to control for a possible momentum or reversal effect.

The monthly illiquidity for stock i is

$$ILL_m^i = (1/D_{im}) \sum_{d=1}^{D_{im}} \sqrt{|R_{md}^i|/VAL_{md}^i} \tag{2}$$

where D_{im} is the number of trading days for stock i in month m.

[5] As of the end of 2010 , the PACAP Japan Database has not generated data beyond 2006 .

All stocks in our sample must meet the following two criteria: 1) It must have valid observations of daily returns and trading values for more than 200 days in a year and more than 15 days in a month so that the illiquidity estimate is more reliable. 2) The month-end stock price must be greater than ¥100 so that stock returns are not affected too much by the minimum tick size of ¥1.[6] Our final sample, ranging from 656 firms in 1975 to 1693 firms in 2006, is described in Table 1.

The monthly market illiquidity is the average illiquidity across stocks in the market portfolio M in month m:

$$ILL_m^M = (1/M) \sum_M ILL_m^i \tag{3}$$

Figure 1a presents the monthly market illiquidity over 1975–2006. It shows that market illiquidity trended down from 1975 to 1990 but increased and then oscillated after that. Correspondingly, as shown in Figure 1b, the Nikkei 225 experienced an upward trend from 1975 to 1989, but declined and oscillated thereafter. According to *Securities Market in Japan* (2001), a publication of the Japan Securities Research Institute, the development of the Japanese securities market from 1975 to 1999 can be divided into several stages: (1) 1975–1984 was the period of coping with the oil crisis; (2) 1985–1989 was the period of the economic bubble; (3) 1990–1999 was the period of financial reform involving debate on and enforcement of Japan's Financial System Reform Law. Roughly, the first two periods coincide with the rapid development of the stock market and a declining trend in illiquidity, while the last period is associated with a market slowdown and an increasing illiquidity trend. Therefore, we divide the whole sample period into two subsample periods in subsequent tests: the first one from 1975 to 1989 and the second one from 1990 to 2006. This also coincides with the sample division of Chang *et al.* (2010), whose further division of the sample period into four stages based on business cycles does not reveal any qualitative difference.

3. The Cross-Sectional Relation between Illiquidity and Stock Returns

In the spirit of Amihud (2002), we specify the following cross-sectional regression model:

$$\begin{aligned} R_m^i = k_{0m} &+ k_{1m} ILLM_{m-1}^i + k_{2m}\beta_{m-1}^i + k_{3m} R_{-Q1}^i \\ &+ k_{4m} R_{-Q2}^i + k_{5m} BM_{m-1}^i + k_{6m} \ln CAP_{m-1}^i \\ &+ k_{7m} STD_{m-1}^i + k_{8m} CP_{m-1}^i + \varepsilon_m^i \end{aligned} \tag{4}$$

where the monthly individual stock return $R^i{}_m$ is a function of illiquidity and a set of control variables. Since the monthly stock illiquidity varies dramatically over time, illiquidity for stock i is further scaled by market illiquidity in month m to obtain the

[6] In addition to tick size, the TSE also has price limit rules for both the maximum daily price changes and maximum price changes between trades. However, the limit between trades is irrelevant, because our illiquidity measure uses daily data. The daily price limits are quite large, ranging from 10% to 30%, and are rarely hit in practice. Therefore, we do not particularly consider them in our sample selection process. Information on tick size schedules and price limit rules for TSE stocks as of 2006 is available from the authors.

Table 1
Sample selection

This table reports the sample selection. The sample period is from July 1975 to December 2006. The stocks included in the sample must have valid observations of daily returns and trading values for more than 200 days in a year and more than 15 days in a month, as well as month-end prices greater than ¥100.

Fiscal year	Trading days	Original stocks in the first section	Stocks with month-end price > ¥100	Stocks with monthly trading days > 15	Stocks with annual trading days > 200	Final sample
1975	286	832	822	740	665	656
1976	286	840	819	771	724	704
1977	285	846	836	791	743	736
1978	285	856	855	771	777	776
1979	286	865	861	756	747	743
1980	285	872	859	761	758	745
1981	285	884	873	769	718	709
1982	286	898	890	795	758	754
1983	287	915	911	806	789	787
1984	284	928	928	859	814	814
1985	285	953	953	908	865	865
1986	274	980	980	957	880	880
1987	274	998	998	970	943	943
1988	261	1023	1023	983	944	944
1989	247	1041	1041	1014	994	994
1990	245	1054	1054	992	967	967
1991	247	1065	1065	962	941	941
1992	247	1077	1077	1026	943	943
1993	247	1094	1094	1063	1013	1013
1994	248	1111	1111	1058	1024	1024
1995	247	1127	1127	1088	1070	1070
1996	247	1171	1171	1132	1083	1083
1997	246	1211	1198	1146	1102	1089
1998	246	1252	1246	1232	1160	1154
1999	247	1321	1287	1303	1238	1204
2000	246	1416	1389	1389	1297	1270
2001	245	1472	1391	1441	1386	1306
2002	247	1524	1500	1517	1453	1430
2003	246	1616	1611	1612	1530	1525
2004	245	1669	1663	1664	1615	1609
2005	247	1696	1689	1695	1667	1660
2006	125	1715	1700	1708	1701	1693

mean-adjusted illiquidity

$$ILLM_m^i = ILL_m^i / ILL_m^M \qquad (5)$$

Other stock characteristics or control variables included in equation (4) are (1) the monthly firm beta, β_{m-1}^i, which is the market beta; (2) the returns of stock i for the past two quarters R_{-Q1}^i and R_{-Q2}^i; (3) the BM, BM_{m-1}^i, which is the most recent yearly book value of equity at the end of year y – 1 (June) divided by the market value of equity

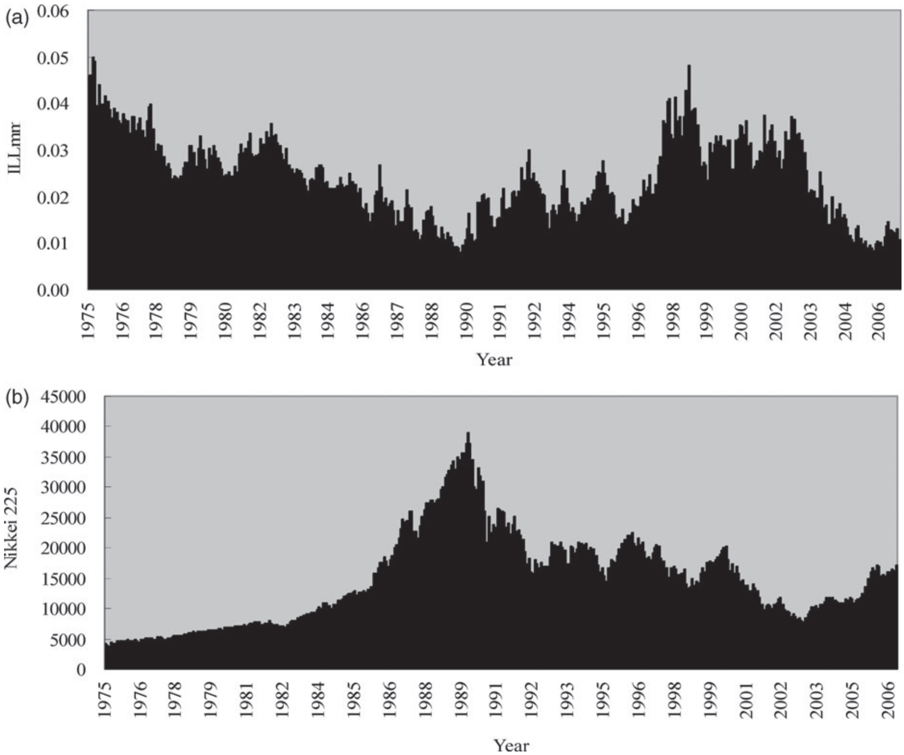

Fig. 1. Time series pattern of monthly market illiquidity and the Nikkei 225

Figures 1a and 1b show the time series pattern of the monthly market illiquidity and the level of the Nikkei 225 during the sample period, respectively, where is calculated as the cross-sectional average of monthly stock illiquidity for all the sample stocks during 1975–2006. The stocks included in the sample must have valid return and trading value data observations for more than 200 days a year and more than 15 days a month. The month-end prices are greater than ¥100. The Nikkei 225 is the index value at the end of each month.

for stock i at the end of the previous month; (4) firm size, $\ln CAP^i_{m-1}$, which is the logarithm of market capitalisation for stock i at the end of month m – 1; (5) total risk, STD^i_{m-1}, which is the standard deviation of the daily return on stock i in month m – 1 (multiplied by 10^2); and (6) cash flow yield (CP^i_{m-1}), which is the ratio of earnings plus depreciation per share in year y – 1 divided by the share price for stock i at the end of month m – 1[7].

Beta is included in equation (4) to control for systematic risk. It is estimated using the Fama–French (1992) methodology. However, stocks are ranked by their market capitalisation each month, instead of each year, and sorted into 25 portfolios. The market model is estimated using daily data for the previous 150 trading days of each portfolio, with the Scholes–Williams (1977) adjustment to obtain the portfolio beta. This portfolio beta is then assigned to each individual stock in the portfolio as its beta

[7] Using annual BM and CP data for each month in a year does not qualitatively affect our results.

risk for that month. The variable STD^i_{m-1} is also included, since investors' portfolios may not be well diversified. Since expected stock returns may be affected by their past returns[8], we include R^i_{-Q1} and R^i_{-Q2} to control for a possible momentum or reversal effect.

We use $\ln CAP^i_{m-1}$ as a control for the well-known size effect. However, as documented in the literature, size itself is often used as a proxy for liquidity. In our case, the Pearson (Spearman) correlation coefficient between our illiquidity measure and size measure is −0.635 (−0.754) for the whole sample period (see Table 2). To avoid multicollinearity, we first make $\ln CAP^i_{m-1}$ orthogonal to illiquidity before including it in our cross-sectional regressions.

Chan *et al.* (1991) document BM^i and CP^i but not dividend yield as major determinants of Japanese stock returns. Hence, our model includes BM^i and CP^i as control variables.

Panel A of Table 2 presents the summary statistics for all the variables described above and the correlation matrix between them. In each month, the cross-sectional mean, standard deviation, skewness, median, minimum, and maximum are calculated for the sample stocks and then averaged over 378 months. The correlations between the variables are presented in panel B of Table 2. Notice that the correlations are generally low except for that between illiquidity and size, as mentioned earlier.

The regression results for equation (4) are reported in Table 3. The first to fourth columns show the results with only four independent variables, namely, illiquidity, firm beta, and two lagged quarterly return terms, in the regression, while the fifth through eighth columns show the results with all eight independent variables included. Following Fama and MacBeth (1973), we estimate equation (4) for each month during our sample period from July 1975 to December 2006 (the first to fifth columns). The mean of the estimated coefficients is calculated for each independent variable, followed by a t-test conducted on the hypothesis that the mean of the coefficients is equal to zero. To control for the famous January effect[9] and to verify whether the cross-sectional relation is stable over time, tests are also performed for the sample excluding January (the second and sixth columns), and for the subsamples 1975–1989 (the third and seventh columns) and 1990–2006 (the fourth and eighth columns).

From the first four columns, we see that the mean of the estimated illiquidity coefficient for all months is 0.0014 and significant at the 5% level. When January is excluded, the coefficient is 0.0013 and also significant at the 5% level. These results are consistent with Amihud's (2002) and suggest the existence of a positive cross-sectional relation between illiquidity and stock returns in general. However, the relation appears unstable over the two subperiods. The mean illiquidity coefficient in the first subsample period is 0.0025, much larger compared to the whole sample result and highly significant at the 1% level, while it is less than 0.0004 for the second subsample period and not significant at all. In contrast, the results reported in Amihud (2002) show that the illiquidity is significant in both subsample periods.[10] Our estimated mean beta coefficient is negative, though not statistically significant, in all four-variable regression samples. However, the insignificant beta coefficient estimates in Japan are consistent with the

[8] Bremer and Hiraki (1999) document that TSE stocks exhibit short-term price reversals.

[9] Hamori (2001) documents the January effect in Japan for the entire period between 1971 and 1997 , although it tends to disappear later in the sample period.

[10] We also tried subsample periods 1975 – 1990 versus 1991 – 2006 and 1975 – 1988 versus 1989 – 2006 . The results are similar.

Table 2

Summary statistics and correlation of stock characteristics

This table presents the summary statistics and correlation matrix for variables used in the cross-sectional regression. In Panel B, the upper triangular matrix reports the Spearman correlation and the lower triangular matrix reports the Pearson correlation. R^i_m is the market return of stock i in month m. The illiquidity measure, ILL^i_m, is the average for month m of the daily ratio of absolute returns to the trading value of stock i in month m. The variable β^i_m is the market beta for stock i in month m as estimated using the Fama–French (1992) method with the Scholes–Williams (1977) adjustment; BM^i_m is the most recent yearly book value of equity at the end of year y – 1 (June) divided by the market value of equity for stock i at the end of month m; CAP^i_m is the market capitalisation of stock i at the end of month m; STD^i_m is the standard deviation of the return for stock i across days in month m; CP^i_m is the ratio of earnings plus depreciation per share in year y – 1 divided by the share price for stock i at the end of month m; R^i_{-Q1} is stock returns over the last three months, and R^i_{-Q2} is the return from month m – 4 to month m – 6. The period covers 1975–2006. The stocks included in the sample must have valid observations of daily returns and trading values for more than 200 days in a year and more than 15 days in a month, and month-end prices greater than ¥100. We multiply most variables by 100 so that we can present the values with only three decimal places.

Panel A: Summary statistics

Variable	Mean of monthly means	Median of monthly means	Minimum of monthly means	Maximum of monthly means	Mean of monthly St.dev.	Mean of monthly skewness
R^i_m (%)	1.143	1.291	−20.726	23.271	9.431	1.827
ILL^i_m (%)	2.444	2.398	0.817	4.980	2.117	1.878
β^i_m	1.058	1.057	0.973	1.172	0.114	0.031
BM^i_m	0.638	0.554	0.244	1.414	0.333	−0.006
CAP^i_m (million ¥)	190,881	206,030	39,419	420,745	584960	10.610
STD^i_m (%)	2.417	2.312	1.735	3.531	0.730	0.874
CP^i_m (%)	6.992	6.142	3.128	11.182	17.865	7.349
R^i_{-Q1} (%)	3.192	4.116	−32.968	39.233	16.484	2.156
R^i_{-Q2}(%)	3.151	4.092	−33.605	39.149	16.434	2.151

Panel B: Correlations

Variable	ILL^i_m	$\ln CAP^i_m$	β^i_m	STD^i_m	CP^i_m	BM^i_m	R^i_{-Q1}	R^i_{-Q2}
ILL^i_m		−0.754	0.111	0.175	0.055	0.242	−0.096	−0.063
$\ln CAP^i_m$	−0.635		−0.166	−0.246	−0.175	−0.337	0.059	0.060
β^i_m	0.011	−0.057		0.130	0.044	−0.004	−0.025	−0.022
STD^i_m	0.194	−0.252	0.072		−0.174	−0.075	−0.020	−0.028
CP^i_m	0.019	−0.015	0.015	−0.064		0.290	−0.042	−0.024
BM^i_m	0.235	−0.342	0.015	0.069	0.035		−0.155	−0.151
R^i_{-Q1}	−0.082	0.040	−0.004	0.065	0.000	−0.138		0.019
R^i_{-Q2}	−0.053	0.039	0.001	0.066	0.004	−0.132	0.019	

finding of Hodoshima *et al.* (2000). On the other hand, R^i_{-Q1} is negative and highly significant while R^i_{-Q2} is not, in all four samples, suggesting short-term price reversals for Japanese stocks. This is inconsistent with Amihud's (2002) finding, that there has been a strong momentum effect for the past 100 days and even the past year. However,

Table 3

Cross-sectional illiquidity effects on stock returns

This table presents the means of the estimated coefficients and the corresponding t-values (in parentheses) from the following monthly cross-sectional regression of stock returns on the respective variables for various samples:

$$R^i_m = k_{0m} + k_{1m}ILLM^i_{m-1} + k_{2m}\beta^i_{m-1} + k_{3m}R^i_{-Q1} + k_{4m}R^i_{-Q2} + k_{5m}BM^i_{m-1} + k_{6m}\ln CAP^i_{m-1} + k_{7m}STD^i_{m-1} + k_{8m}CP^i_{m-1} + \varepsilon^i_m$$

where $ILLM^i_{m-1}$ is the mean-adjusted illiquidity for stock i across the days in month m – 1; β^i_{m-1} is the market beta for stock i in month m – 1, which is estimated using the Fama–French (1992) method with the Scholes–Williams (1977) adjustment; R^i_{-Q1} is the stock return over the last three months and R^i_{-Q2} is the return from month m – 4 to month m – 6; BM^i_{m-1} is the most recent yearly book value of equity at the end of year y – 1 divided by the market value of equity for stock i at the end of month m – 1; $\ln CAP^i_{m-1}$ is the logarithm of the market capitalisation for stock i at the end of month m – 1 after being made orthogonal to $ILLM^i_{m-1}$; STD^i_{m-1} is the standard deviation of returns for stock i across the days of month m – 1 (multiplied by 10^2); and CP^i_{m-1} is the ratio of earnings plus depreciation per share in year y – 1 divided by the share price for stock i at the end of month m – 1. The monthly returns are from July 1975 to December 2006, and the stock characteristics are from June 1975 to November 2006.**, and * denote significance at the 1% and 5% levels, respectively.

Variable	All months	Excl. January	1975–1989	1990–2006	All months	Excl. January	1975–1989	1990–2006
Constant	0.0105	0.0124	0.0231**	−0.0010	0.0142*	0.0184**	0.0310**	−0.0013
	(1.652)	(1.879)	(2.827)	(−0.104)	(2.215)	(2.764)	(3.817)	(−0.137)
$ILLM^i_{m-1}$	0.0014*	0.0013*	0.0025**	0.0004	0.0019**	0.0016*	0.0026**	0.0006
	(2.311)	(2.162)	(2.678)	(0.504)	(2.855)	(2.509)	(2.834)	(0.741)
β^i_{m-1}	−0.0051	−0.0084	−0.0087	−0.0019	−0.0034	−0.0060	−0.0097	0.0024
	(−0.791)	(−1.242)	(−1.258)	(−0.179)	(−0.604)	(−1.010)	(−1.488)	(0.270)
R^i_{-Q1}	−0.0259**	−0.0196**	−0.0237**	−0.0278**	−0.0276**	−0.0212**	−0.0269**	−0.0283**
	(−4.479)	(−3.414)	(−3.715)	(−2.963)	(−4.909)	(−3.756)	(−4.010)	(−3.180)
R^i_{-Q2}	−0.0008	0.0038	0.0026	−0.0038	−0.0001	0.0054	0.0034	−0.0032
	(−0.189)	(0.930)	(0.470)	(−0.678)	(−0.016)	(1.313)	(0.553)	(−0.601)
BM^i_{m-1}					0.0181**	0.0179**	0.0254**	0.0113**
					(6.415)	(5.946)	(4.294)	(3.737)
$\ln CAP^i_{m-1}$					−0.0013	−0.0006	−0.0022	−0.0004
					(−1.392)	(−0.887)	(−1.522)	(−0.531)
STD^i_{m-1}					−0.0018	−0.0033**	−0.0015	−0.0020
					(−1.530)	(−2.879)	(−1.032)	(−1.138)
CP^i_{m-1}					−0.0138**	−0.0123**	−0.0332**	0.0041
					(−5.419)	(−4.624)	(−7.585)	(1.942)

it is consistent with Bremer and Hiraki (1999) that TSE stocks exhibit short-term price reversals[11].

When all eight independent variables are included, the results are largely the same, in the sense that the illiquidity coefficient is still positively and significantly related to the stock returns in the whole sample, the sample with January excluded, and the first subsample period, but it is close to zero and insignificant in the second subsample period. The estimated beta coefficient is insignificant and mostly negative, and the lagged stock return R^i_{-Q1} has a negative and significant impact on the current stock return in all samples, while R^i_{-Q2} does not.

For additional control variables, we find that coefficient estimates for $\ln CAP^i_{m-1}$ are consistently insignificant and mostly negative in the fifth to eighth columns of Table 3. This may be due to the fact that the $\ln CAP^i_{m-1}$ included in the regression is only the residual part that is orthogonal to our illiquidity measure. The estimated coefficient for STD^i_{m-1} is negative but mostly insignificant. The variable BM^i_{m-1} has a positive and significant impact on stock returns in all four samples, which is consistent with Chan *et al.* (1991). In addition, CP^i_{m-1} has a significant impact on stock returns; however, its sign is not consistent across samples: It is negative for the whole sample, the sample excluding January, and the first subsample period, but it is positive in the second subsample period, again indicating something unique about the second subsample period.

On the whole, our results suggest that illiquidity is priced in the Japanese market, but not in the second subsample period. This pattern is robust across different specifications. In fact, Chang *et al.* (2010) also find that the Amihud measure is unrelated to stock returns during the period 1990–2004. From Figure 1, we see that the first subsample period corresponds to a booming market trend with declining illiquidity, while the second subsample period coincides with a down and oscillating market and a trend of increasing and fluctuating illiquidity. Further examination of the monthly market excess returns shows that, of the 180 months in the first subsample period, 115 are associated with positive market excess returns and 65 with negative market excess returns. Of the 198 months in the second subsample period, only 97 are associated with positive market excess returns, and 101 with negative market excess returns. Therefore, the ratio of negative excess market return months to positive ones is much higher in the second subsample period (101/97) than in the first period (65/115).

Using Japanese stock return data from 1956 to 1995, Hodoshima *et al.* (2000) find that regressing returns on beta without differentiating between positive and negative market excess returns leads to a flat relation between beta and returns. However, significant conditional positive or negative relations between beta and returns are found once the observations are separated into positive and negative market premium groups, where the positive and negative market premiums refer to $R_{m-} R_f > 0$ and $R_{m-} R_f < 0$, respectively.[12] The authors explain that the expected market excess return should

[11] We also follow Amihud (2002) to include past 100-day returns and the returns of the past year minus the past 100-day returns in the regression. The results are largely the same; that is, we find negative and significant coefficients for the past 100-day returns, but insignificant coefficients for further past returns.

[12] Chan and Lakonishok (1993), Grundy and Malkiel (1996), and Pettengill *et al.* (1995) investigate the relation between returns and beta by considering whether the market excess return is positive or negative in the US market. These authors find that beta and stock returns are significantly related.

never be negative, but actual observations used in the regression often are. Similarly, one can argue that the expected illiquidity premium should never be negative, but the realised premium may well be so. If the realised illiquidity premium is positively correlated with excess market returns, then the estimated relation between the Amihud measure and stock returns may be distorted. Therefore, we repeat the cross-sectional regressions by discarding all months associated with the negative market premium observations to see if illiquidity and beta are priced, especially in the second subsample period.

Table 4 presents the estimation results. Consistent with Hodoshima *et al.* (2000), beta is mostly positive and significant. However, estimated coefficients for illiquidity are still qualitatively the same as those reported in Table 3. The coefficient of illiquidity remains insignificant in the second subsample period, whether for four- or eight-variable regressions. This suggests that the insignificant mean estimate for the illiquidity coefficient in the second subsample period is not due to the concentration of more negative excess market returns in the period.

The second subsample period may be special because it coincided with many changes in the Japanese financial system. Hamao *et al.* (2003) document that idiosyncratic volatility in Japanese stocks has fallen, coinciding with a slowdown in the capital allocation process within the Japanese economy. The authors opine that Japanese corporate managers may have chosen to bail out large companies in their respective keiretsu rather than allocate capital to young companies, which caused the stock prices of Japanese stocks in the 1990s to be more correlated than those of the 1980s. It is possible that corporate behaviour as well as investor risk tolerance may have undergone some big changes in the 1990s, such that illiquidity risk was not captured by the Japanese data during this period.

We also repeat the regressions for the months when the market premium is negative. As expected, beta is mostly negative and significant, but the estimated illiquidity coefficient remains insignificant in the second subsample period. We further perform a few more robustness checks: (1) We sort the portfolios annually rather than monthly for beta estimation and use daily returns in the previous year to estimate the beta; (2) we use annual data to compute BM and CP, as in Chan *et al.* (1991); and (3) we use an individual firm illiquidity measure without scaling with market illiquidity. All these factors do not qualitatively affect our results, which are not reported here to save space.

4. Time Series Relation between Illiquidity and Stock Returns

Amihud (2002) argues that stocks are not only riskier but also less liquid than short-term treasury securities. Hence, stock returns in excess of the T-bill rate (risk premium) include a premium for illiquidity. It follows that if investors anticipate higher market illiquidity, they will expect higher returns. More specifically, expected stock returns should be positively related to expected illiquidity, while unexpected illiquidity should be negatively related to contemporaneous stock returns.

We use the following AR(1) model to generate the expected and unexpected market illiquidity:

$$ILL_m^M P_{m-1}^M = c_0 + c_1 ILL_{m-1}^M P_{m-1}^M + v_m \qquad (6)$$

Table 4

Cross-sectional illiquidity effects on stock returns for months with positive market premium

This table presents the means of the estimated coefficients and corresponding t-values (in parentheses) from the following monthly cross-sectional regression of stock returns on the respective variables for various samples:

$$R^i_m = k_{0m} + k_{1m} ILLM^i_{m-1} + k_{2m}\beta^i_{m-1} + k_{3m} R^i_{-1Q} + k_{4m} R^i_{-Q2} + k_{5m} BM^i_{m-1} + k_{6m} \ln CAP^i_{m-1} + k_{7m} STD^i_{m-1} + k_{8m} CP^i_{m-1} + \varepsilon^i_m$$

Only months with a positive market premium are included in the regression. Here, $ILLM^i_{m-1}$is the mean-adjusted illiquidity for stock i across the days in month $m-1$; β^i_{m-1} is the market beta for stock i in month $m-1$ as estimated using the Fama–French (1992) method with the Scholes–Williams (1977) adjustment; R^i_{-Q1} is stock returns over the last three months and R^i_{-Q2} is the return from months $m-4$ to $m-6$; BM^i_{m-1} is the most recent yearly book value of equity at the end of year $y-1$ (June) divided by the market value of equity for stock i at the end of month $m-1$; $\ln CAP^i_{m-1}$ is the logarithm of the market capitalisation for stock i at the end of month $m-1$ after made orthogonal to$ILLM^i_{m-1}$; STD^i_{m-1} is the standard deviation of returns for stock i across the days of month $m-1$ (multiplied by 10^2); and CP^i_{m-1} is the ratio of earnings plus depreciation per share in year $y-1$ divided by the share price for stock i at the end of month $m-1$. The monthly returns are from July 1975 to December 2006, and the stock characteristics are from June 1975 to November 2006.**, and* denote significance at the 1% and 5% levels, respectively.

Variable	All months	Excl. January	1975–1989	1990–2006	All months	Excl. January	1975–1989	1990–2006
Constant	0.0159	0.0178	0.0426**	−0.0138	0.0068	0.0106	0.0335**	−0.0210
	(1.817)	(1.909)	(4.843)	(−0.910)	(0.797)	(1.186)	(3.505)	(−1.516)
$ILLM^i_{m-1}$	0.0037**	0.0034**	0.0061**	0.0012	0.0027**	0.0028**	0.0053**	−0.0001
	(3.917)	(3.307)	(4.530)	(0.874)	(2.738)	(2.677)	(3.960)	(−0.076)
β^i_{m-1}	0.0217*	0.0192*	−0.0094	0.0564**	0.0175**	0.0159*	0.0060	0.0295**
	(2.594)	(2.176)	(−1.326)	(3.727)	(3.096)	(2.585)	(1.020)	(3.050)
R^i_{-Q1}	−0.0190**	−0.0190**	−0.0045	−0.0352**	−0.0188**	−0.0182**	−0.0063	−0.0317**
	(−4.457)	(−4.669)	(−1.043)	(−4.819)	(−4.834)	(−4.743)	(−1.419)	(−5.145)
R^i_{-Q2}	−0.0028	−0.0044	0.0127	−0.0201**	−0.0053	−0.0054	0.0083	−0.0195**
	−(0.533)	(−0.794)	(1.764)	(−2.696)	(−1.027)	(−0.996)	(1.102)	(−2.880)
BM^i_{m-1}					0.0038	0.0037	0.0045	0.0030
					(1.195)	(1.073)	(0.893)	(0.804)
$\ln CAP^i_{m-1}$					−0.0031**	−0.0027**	−0.0034*	−0.0027*
					(−3.442)	(−2.797)	(−2.384)	(−2.565)
STD^i_{m-1}					0.0084**	0.0068**	0.0040*	0.0130**
					(6.260)	(5.157)	(2.493)	(6.255)
CP^i_{m-1}					−0.0013	−0.0003	0.0013	−0.0040
					(−0.358)	(−0.064)	(0.191)	(−1.213)

where ILL_m^M is the monthly market illiquidity and v_m is the residual representing the unexpected market illiquidity$ILLU_m^M$. We do not use logarithmic transformations for illiquidity, because its square root is already transformed, as described in the previous section. The variable P_{m-1}^M is the ratio of the capitalisation of the market portfolio at the end of month $t-1$ to that at the end of June 1975. A potential problem with using the Amihud illiquidity measure in the time series regression is that it is measured in percent per ¥1 million. As pointed out by Acharya and Pedersen (2005), the Amihud measure may be non-stationary (*e.g.*, inflation is ignored); however, we want to relate it to stock returns, which are stationary. This is not an issue in cross-sectional regressions, but it is in time series regressions. Multiplying ILL_m^M with P_{m-1}^M can make the illiquidity measure relatively stationary. Following Acharya and Pederson (2005), we use the same month of the market index (P_{m-1}^M) to adjust for the illiquidity measure in months m and $m-1$. This is to ensure that we are measuring expected and unexpected illiquidity rather than the change in market index.

Let us define adjusted illiquidity in month m as $AILL_m^M = ILL_m^M P_{t-1}^M$, and that in month $m-1$ as $AILL_{m-1}^M = ILL_{m-1}^M P_{t-1}^M$. The expected illiquidity after adjustment $AILLE_m^M$ in month m is determined by

$$AILLE_m^M = c_0 + c_1 AILL_{m-1}^M \tag{7}$$

Then the expected market excess return is generated by

$$R_m^M - R_m^f = f_0 + f_1 AILLE_m^M + u_m = g_0 + g_1 AILL_{m-1}^M + u_m \tag{8}$$

where $g_0 = f_0 + f_1 c_0$, $g_1 = f_1 c_1$, $R_m^M$is the return of the market portfolio M (all stocks in our sample) in month m,[13] and R_m^f is the monthly call money rate, or one-month Gensaki rate, for month m, as in Chan *et al.* (1991), which can be retrieved from the monthly key economic file in the PACAP Database. The term u_m can be decomposed into the unexpected illiquidity $ILLU_m^M$ (or $AILLU_m^M$) and an error term w_m. After controlling for the January effect, the time series regression of excess market returns on market illiquidity is[14]

$$R_m^M - R_m^f = g_0 + g_1 AILL_{m-1}^M + g_2 AILLU_m^M + g_3 JAN_m + w_m \tag{9}$$

The two testable hypotheses are as follows.

Hypothesis 1: Expected market illiquidity is positively related to expected market excess return, that is, $g_1 > 0$.
Hypothesis 2: Unexpected market illiquidity is negatively related to contemporaneous market excess returns, that is, $g_2 < 0$.

Amihud (2002) further proposes and tests the flight to liquidity hypothesis. Increases in expected market illiquidity have two effects on expected stock returns: On the one hand, the stock price declines and expected returns rise for all stocks; on the other hand, capital flies from less liquid to more liquid stocks. These two effects reinforce each other

[13] It is calculated as the log difference of two consecutive month-end daily closing prices.

[14] A similar specification has been used by French *et al.* (1987) in testing the effect of risk on excess stock returns.

for illiquid stocks, but offset each other for liquid ones. Increasing market illiquidity not only negatively affects prices for illiquid stocks but also induces investors to switch to more liquid stocks, which further depresses the price for illiquid stocks. However, increasing market illiquidity leads to an increase in demand for liquid stocks, which mitigates their price decline. Therefore, the illiquidity effect should be stronger for small stocks and weaker for large stocks because firm size is negatively correlated with illiquidity. Replacing market portfolio return series with size portfolio return series,[15] we can rewrite equation (9) as

$$R_m^P - R_m^f = g_0^P + g_1^P AILL_{m-1}^P + g_2^P AILLU_m^P + g_3^P JAN_m + w_m^P \tag{10}$$

where p denotes one of the 10 size-based portfolios (portfolio 10 contains the largest stocks) and R_m^p is the average return across stocks in portfolio p for month m. The testable hypotheses are as follows.

Hypothesis 3: The expected illiquidity effect g_1^p *should be positive but decrease as portfolio size increases, that is,* $g_1^2 > g_1^4 > g_1^6 > g_1^8 > g_1^{10} > 0$.
Hypothesis 4: The unexpected illiquidity effect g_2^p *should be negative but increase as portfolio size increases, that is,* $g_2^2 < g_2^4 < g_2^6 < g_2^8 < g_2^{10} < 0$.

Table 5 presents the estimation results for the whole sample period from July 1975 to December 2006 (378 months). Several observations are evident. First, expected illiquidity $AILL_{m-1}^p$ has a positive and significant impact on the expected return for the whole market, as well as for all portfolios; however, there is no monotonically decreasing pattern for the estimated coefficients across size portfolios. This is consistent with H1, that expected illiquidity should be positively and significantly related to expected stock returns, but inconsistent with H3, that expected illiquidity effect should decrease with firm size. Second, unexpected illiquidity $AILLU_m^p$ is negatively associated with expected stock returns, and the estimated coefficient is significant and monotonically increasing with firm size. This finding is consistent with both H2, that contemporaneous stock returns are negatively associated with unexpected illiquidity, and H4, that the unexpected illiquidity effect is negative and increases with firm size. Finally, we find that the January effect is more significant for smaller stock portfolios but insignificant for the portfolio with the largest firms. These results are consistent with the findings in Amihud (2002), except we do not find a monotonically decreasing coefficient for expected illiquidity with increasing firm size.

Table 6 further presents the results for the two subsample periods. These results are consistent across the two subsample periods and also consistent with the results in Table 5. However, it is worth mentioning that the estimated coefficient for expected illiquidity is much larger in the first subsample period than in the second subsample period, whereas for unexpected illiquidity, it is much smaller in the first subsample period (more negative). This suggests that expected illiquidity is priced less while unexpected illiquidity is priced more in the Japanese stock market in the second subsample period than in the first subsample period. This finding may be related with that in Section 3, that illiquidity levels are priced in the first subsample period, but not in the second. The illiquidity measure used in the cross-sectional regressions is the lagged level of illiquidity, which is similar to the expected illiquidity used in

[15] The 10 size portfolios are the same as those formed in Section 3.

Table 5

Time series illiquidity effects on stock returns

This table presents time series regression results for the market, as well as size-sorted portfolios. The regression equation is

$$R_m^P - R_m^f = g_0^P + g_1^P AILL_{m-1}^P + g_2^P AILLU_m^P + g_3^P JAN_m + w_m^P$$

where R_m^p is the monthly equally weighted size portfolio return and $p = 2, 4, 6, 8$, and 10. Of course, when $p = M$, it is the equally weighted market return. Here, R_m^f is the one-month Gensaki monthly rate, $AILL_{m-1}^p$ is the expected monthly portfolio illiquidity, $AILLU_m^p$ is the unexpected monthly portfolio illiquidity, and JAN_m is a January dummy that equals one in January and zero otherwise. The sample period is from July 1975 to December 2006. **, and * denote significance at the 1% and 5% levels, respectively.

Portfolio	Constant	$AILL_{m-1}$	$AILLU_m$	JAN_m	R^2
Market	0.075**	0.022**	−0.084**	0.026*	0.161
	(3.767)	(5.275)	(−5.935)	(2.340)	(0.154)
Portfolio 2	0.060*	0.020**	−0.132**	0.048**	0.197
	(2.477)	(3.908)	(−7.677)	(3.512)	(0.190)
Portfolio 4	0.072**	0.022**	−0.107**	0.038**	0.175
	(3.207)	(4.636)	(−6.687)	(2.996)	(0.168)
Portfolio 6	0.074**	0.021**	−0.091**	0.031**	0.157
	(3.450)	(4.867)	(−5.987)	(2.606)	(0.150)
Portfolio 8	0.082**	0.023**	−0.071**	0.018	0.142
	(4.153)	(5.639)	(−5.025)	(1.586)	(0.135)
Portfolio 10	0.081**	0.022**	−0.052**	0.013	0.118
	(4.271)	(5.677)	(−3.837)	(1.207)	(0.111)

this section, albeit the former is for individual stocks while the latter is for the whole market or size portfolios. However, we are unable to fully explain why liquidity is priced less in the second subsample period, which deserves further exploration in future research.

We also perform the following robustness checks. First, Kendall (1954) points out that the estimated coefficient $\hat{c}_1$ from the finite samples is biased downward in AR(1) models such as equation (7). The author proposes a simple but accurate bias correction approximation procedure: Augment the estimated coefficient $\hat{c}_1$ by the term $(1 + 3\hat{c}_1)/T$, where T is the sample size. We apply Kendall's bias correction in the computation of unexpected illiquidity. Second, we include the term yield premium in the estimation.[16] Third, we repeat the same tests using portfolios formed on the BM rather than market capitalisation. None of these checks qualitatively change the results reported in Tables 5 and 6. The results are available upon request.

[16] The term yield premium $TM_m = YL_m - R_m^{G3}$ is computed as the difference between the yield to maturity of 10-year government bonds (YL_m) and the three-month Gensaki rate in month m (R_m^{G3}). The default yield premium is not included in the regression because we do not have the data for Japan.

Table 6

Time series illiquidity effects on stock returns for the two subsample periods

This table presents time series regression results for the market, as well as size-sorted portfolios. The regression equation is

$$R_m^P - R_m^f = g_0^P + g_1^P AILL_{m-1}^P + g_2^P AILLU_m^P + g_3^P JAN_m + w_m^P$$

where R_m^p is the monthly equally weighted size portfolio return and $p = 2, 4, 6, 8$, and 10. Of course, when $p = M$, it is the equally weighted market return. Here, R_m^f is the one-month Gensaki monthly rate, $AILL_{m-1}^p$ is the expected monthly portfolio illiquidity, $AILLU_m^p$ is the unexpected monthly portfolio illiquidity, and JAN_m is a January dummy that equals one in January and zero otherwise. Panel A reports the results for the first subsample period, July 1975 to December 1989, and Panel B presents the results for the second subsample period, January 1990 to December 2006.**, and * denote significance at the 1% and 5% levels, respectively.

Portfolio	Constant	$AILL_{m-1}$	$AILLU_m$	JAN_m	R^2
Panel A: 1975–1989					
Market	0.086	0.026*	−0.064**	0.035**	0.168
	(1.562)	(2.424)	(−3.616)	(3.045)	(0.154)
Portfolio 2	0.070	0.023	−0.108**	0.058**	0.245
	(1.050)	(1.794)	(−5.097)	(4.248)	(0.233)
Portfolio 4	0.085	0.026*	−0.084**	0.045**	0.192
	(1.340)	(2.112)	(−4.155)	(3.465)	(0.178)
Portfolio 6	0.094	0.027*	−0.071**	0.041**	0.182
	(1.581)	(2.398)	(−3.738)	(3.409)	(0.168)
Portfolio 8	0.113	0.031**	−0.053**	0.027*	0.129
	(1.914)	(2.738)	(−2.809)	(2.249)	(0.115)
Portfolio 10	0.095	0.027*	−0.032	0.018	0.074
	(1.625)	(2.397)	(−1.712)	(1.542)	(0.058)
Panel B: 1990–2006					
Market	0.033	0.010	−0.101**	0.020	0.130
	(1.148)	(1.535)	(−4.831)	(1.092)	(0.116)
Portfolio 2	0.029	0.011	−0.149**	0.040	0.175
	(0.804)	(1.325)	(−5.774)	(1.739)	(0.161)
Portfolio 4	0.033	0.011	−0.124**	0.033	0.152
	(1.027)	(1.459)	(−5.262)	(1.572)	(0.139)
Portfolio 6	0.032	0.009	−0.095**	0.019	0.115
	(1.080)	(1.436)	(−4.502)	(1.005)	(0.101)
Portfolio 8	0.034	0.010	−0.086**	0.011	0.105
	(1.227)	(1.553)	(−4.290)	(0.601)	(0.091)
Portfolio 10	0.036	0.010	−0.069**	0.010	0.082
	(1.371)	(1.647)	(−3.615)	(0.575)	(0.068)

5. Liquidity Risk and Stock Returns

Following Acharya and Pedersen (2005), this section tests if liquidity risk is priced over and above the illiquidity level and the market risk in Japan. We define a one-beta CAPM for net returns as

$$E_t\left(R_{t+1}^i - C_{t+1}^i\right) = R^f + \lambda_t \frac{\mathrm{cov}_t\left(R_{t+1}^i - C_{t+1}^i, R_{t+1}^M - C_{t+1}^M\right)}{\mathrm{var}_t\left(R_{t+1}^M - C_{t+1}^M\right)} \tag{11}$$

where $R_t^i = \frac{d_t^i + p_t^i}{p_{t-1}^i}$ is the expected gross return of an asset; $C_t^i = \frac{c_t^i}{p_{t-1}^i}$ is the relative illiquidity cost; $R_t^M = \frac{\sum_i s^i (d_t^i + p_t^i)}{\sum_i s^i p_{t-1}^i}$ is the market return; $C_t^M = \frac{\sum_i s^i c_t^i}{\sum_i s^i p_{t-1}^i}$ is the relative market illiquidity; d_t^i and c_t^i refer to the dividend paid and illiquidity cost (the cost of selling the asset) incurred for security i during each time t, respectively; p_t^i denotes the price for security i at the end of time t; s^i denotes the number of shares for security i; and $\lambda = E(R_t^M - C_t^M - R^f)$ is the risk premium after the illiquidity cost. Further expanding and rewriting equation (11), we obtain the following liquidity-adjusted CAPM:[17]

$$E_t\left(R_{t+1}^i\right) = R^f + E_t\left(C_{t+1}^i\right) + \lambda_t \frac{\mathrm{cov}_t\left(R_{t+1}^i, R_{t+1}^M\right)}{\mathrm{var}_t\left(R_{t+1}^M - C_{t+1}^M\right)} + \lambda_t \frac{\mathrm{cov}_t\left(C_{t+1}^i, C_{t+1}^M\right)}{\mathrm{var}_t\left(R_{t+1}^M - C_{t+1}^M\right)} - \lambda_t \frac{\mathrm{cov}_t\left(R_{t+1}^i, C_{t+1}^M\right)}{\mathrm{var}_t\left(R_{t+1}^M - C_{t+1}^M\right)} - \lambda_t \frac{\mathrm{cov}_t\left(C_{t+1}^i, R_{t+1}^M\right)}{\mathrm{var}_t\left(R_{t+1}^M - C_{t+1}^M\right)} \tag{12}$$

Assuming constant conditional covariances of innovations in illiquidity and returns, we can express equation (12) as

$$E\left(R_t^i - R_t^f\right) = E\left(C_t^i\right) + \lambda\beta^{1i} + \lambda\beta^{2i} - \lambda\beta^{3i} - \lambda\beta^{4i} \tag{13}$$

This simple theoretical framework suggests that (1) the liquidity risk affects asset prices through the last three terms in the equation and (2) the effects of the liquidity level $E_t(C_{t+1}^i)$ and the liquidity risk are separate. The market risk is captured by β^{1p}. The three liquidity risk factors are commonality in liquidity, β^{2p}, individual stock return sensitivity to market illiquidity, β^{3p}, and individual stock illiquidity sensitivity to market return, β^{4p}.

Commonality in liquidity,$\mathrm{cov}_{t-1}(C_t^i, C_t^M)$, should be positively associated with expected returns because investors need to be compensated for the risk that their assets may become illiquid when the market as a whole is illiquid. Individual return sensitivity to market illiquidity, $\mathrm{cov}_{t-1}(R_t^i, C_t^M)$, should have a negative impact on expected returns because investors are willing to accept lower returns on assets normally associated with high returns in times of market illiquidity. Individual illiquidity sensitivity to market returns,$\mathrm{cov}_{t-1}(C_t^i, R_t^M)$, should also have a negative impact on expected returns because investors are willing to accept lower returns on assets with higher liquidity when market returns are low.

Using Center for Research in Security Prices (CRSP) data from 1962 to 1999, Acharya and Pedersen (2005) show that the required return of a security i does increase in the covariance between its illiquidity and market illiquidity,$\mathrm{cov}_{t-1}(C_t^i, C_t^M)$, and decreases in the covariance between the security's return and market illiquidity, $\mathrm{cov}_{t-1}(R_t^i, C_t^M)$, as well as its illiquidity and market return, $\mathrm{cov}_{t-1}(C_t^i, R_t^M)$. In addition, the authors show that the liquidity-adjusted CAPM explains the data better than the standard CAPM. Furthermore, they present evidence that liquidity risk is important over and above the effects of market risk and the level of liquidity.

To estimate illiquidity costs and various betas, we first compute the return and normalised illiquidity cost for each firm each month. The normalised illiquidity cost of selling over the ex-dividend price (c_t^i/p_{t-1}^i) is proxied by the transformation of Amihud's measure, as suggested by Acharya and Pedersen (2005). Second, we form a market portfolio and 25 test portfolios sorted on the basis of illiquidity or BM and compute their

[17] For the model's assumptions and derivation details, see Acharya and Pedersen (2005) .

monthly returns and normalised illiquidity. Third, we estimate illiquidity innovations, and thus liquidity betas, for each portfolio. We then test the liquidity-adjusted CAPM by applying the generalised method of moments (GMM) to estimate equation (13) and its variations.[18]

Table 7 reports the estimated betas and other properties for odd-numbered illiquidity-sorted portfolios: Panel A presents the results for the whole sample period 1975–2006,[19] while panels B and C present the results for the first (1975–1989) and second (1990–2006) subsample periods, respectively.

We see in panel A of Table 7 that a sort on past illiquidity produces portfolios with monotonically increasing average illiquidity ($E(C^p)$) and the commonality in illiquidity beta (β^{2p}) from portfolios 1 through 25[20] but with a largely decreasing portfolio return sensitivity to market illiquidity (β^{3p}) and illiquidity sensitivity to market returns (β^{4p}) from portfolios 1 through 25. All portfolio betas have predicted signs with significant t-values. According to Acharya and Pedersen (2005), such an association between the illiquidity level and the liquidity betas supports the flight to liquidity hypothesis. That is, the higher the illiquidity level, the higher the illiquidity risk.

In addition, we see that stocks with higher average illiquidity tend to have a higher market beta (β^{1p}), a higher volatility of illiquidity innovation ($\sigma(\Delta c^p)$), higher BM, a higher volatility of portfolio returns ($\sigma(R^p)$), smaller market capitalisation (*CAP*), and lower trading turnover. More importantly, we find that the portfolio excess return $E(R^p - R^f)$ exhibits an increasing, though not monotonically, trend with the illiquidity-sorted portfolios. All these findings are very consistent with theoretical predictions, as well as the empirical findings in Acharya and Pedersen (2005).

The subsample results reported in panels B and C of Table 7 are qualitatively the same, except that portfolio excess returns in the second subsample period do not show a clear trend with illiquidity-sorted portfolios. This echoes the finding in previous sections that the liquidity effect in the second subsample period is weaker in the Japanese stock market.

We also compute similar statistics for odd-numbered portfolios sorted by a standard deviation of the illiquidity. The results are generally the same as those presented in Table 7. We further compute the betas for portfolios sorted by BM. Again the results are largely the same. These results are available upon request.

The correlation matrices for the betas of illiquidity-sorted portfolios in the whole sample period as well as subsample periods are shown in Table 8. We find high and significant correlation coefficients between these betas, as shown in Table 8. This is not surprising, given the results in Table 7, but it does suggest that we face a multicollinearity problem when estimating equation (13).[21]

Table 9 presents the GMM estimates of equation (13) and its variations. Panel A of Table 9 reports the regressions for illiquidity-sorted portfolios for the whole sample

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[18] The GMM point estimates are the same as those derived using ordinary least squares, either in a pooled regression or using the Fama–MacBeth (1973) method. However, the GMM is more efficient because it takes into account possible heteroskedasticity and serial correlation.

[19] We use the entire time series to compute the four betas, as Black *et al.* (1990) and Pastor and Stambaugh (2003) do. Therefore, they are not directly comparable to the yearly betas computed in Section 3.

[20] Portfolio 1 is the least illiquid, while portfolio 25 is the most illiquid.

[21] Acharya and Pedersen (2005) find even higher correlation coefficients between these four betas.

Table 7

Portfolio properties

This table reports the properties of the odd-numbered illiquidity-sorted portfolios formed each year during 1975–2006. The market beta (β^{1p}) and the liquidity betas (β^{2p},β^{3p}, and β^{4p}) are computed using the monthly returns and illiquidity observations of each test portfolio and the equally weighted market portfolio for the whole period as well as for the two subperiods. The standard deviation of illiquidity innovations is reported in the column for $\sigma(\Delta c^p)$. The average illiquidity $E(C^p)$, the average excess return $E(R^p - R^f)$, the average turnover (*TRN*), the average market capitalisation (*CAP*), and the average book-to-market ratio (*BM*) are also computed for each portfolio as time series averages of the respective monthly characteristics. Here $\sigma(R^p)$ is the average of the standard deviation of daily returns for the portfolio's constituent stocks computed each month. The *t*-statistics are reported in parentheses. Since the magnitudes of the estimated β^{2p},β^{3p}, and β^{4p} are small, we multiply all estimated betas by 100.

Panel A: 1975–2006

Portfolio	β^{1p} (·100)	β^{2p} (·100)	β^{3p} (·100)	β^{4p} (·100)	$E(C^p)$ (%)	$\sigma(\Delta c^p)$ (%)	$E(R^p - R^f)$ (%)	$\sigma(R^p)$ (%)	*TRN* (%)	*CAP* (billion ¥)	*BM*
1	59.38 (14.25)	0.0000 (6.84)	−0.44 (−5.29)	0.00 (−9.39)	0.25	0.00	0.54	1.94	6.33	2051.39	0.41
3	79.50 (27.16)	0.0003 (10.25)	−0.49 (−6.12)	−0.02 (−10.29)	0.26	0.00	0.53	2.07	6.74	447.53	0.47
5	87.39 (36.33)	0.0007 (14.11)	−0.60 (−7.46)	−0.04 (−10.64)	0.26	0.00	0.46	2.16	6.47	236.86	0.50
7	91.85 (41.93)	0.0014 (16.26)	−0.59 (−7.18)	−0.07 (−11.43)	0.27	0.01	0.49	2.20	5.79	164.34	0.57
9	99.33 (47.10)	0.0026 (20.31)	−0.65 (−7.40)	−0.11 (−11.89)	0.29	0.01	0.67	2.26	6.02	124.90	0.59
11	102.00 (53.95)	0.0043 (20.24)	−0.68 (−7.79)	−0.16 (−9.82)	0.31	0.02	0.37	2.25	5.44	89.51	0.62
13	105.63 (52.76)	0.0075 (25.46)	−0.77 (−8.61)	−0.28 (−10.62)	0.34	0.03	0.59	2.28	5.45	72.64	0.67
15	102.67 (53.75)	0.0113 (24.43)	−0.77 (−8.94)	−0.37 (−9.05)	0.39	0.05	0.72	2.30	5.03	62.92	0.69
17	101.95 (54.11)	0.0179 (28.87)	−0.74 (−8.56)	−0.52 (−8.35)	0.46	0.07	0.56	2.28	4.07	50.77	0.73
19	105.40 (51.44)	0.0262 (32.53)	−0.75 (−8.37)	−0.82 (−9.51)	0.55	0.10	0.69	2.32	3.98	39.84	0.76
21	99.61 (47.75)	0.0451 (41.11)	−0.77 (−9.10)	−1.08 (− 7.31)	0.69	0.17	0.87	2.37	3.71	35.90	0.78
23	107.78 (46.27)	0.0701 (36.29)	−0.79 (−8.49)	−1.64 (−6.92)	0.99	0.27	0.97	2.54	3.35	28.67	0.80
25	116.29 (39.30)	0.2394 (45.37)	−0.94 (−9.21)	−5.52 (−7.13)	2.23	0.89	1.27	2.90	2.52	16.60	0.78

Table 7
Continued

Portfolio	β^{1p} (·100)	β^{2p} (·100)	β^{3p} (·100)	β^{4p} (·100)	$E(C^p)$ (%)	$\sigma(\Delta c^p)$ (%)	$E(R^p - R^f)$ (%)	$\sigma(R^p)$ (%)	*TRN* (%)	*CAP* (billion ¥)	*BM*
Panel B:	1975–1989										
1	60.97 (5.89)	0.0000 (7.05)	−0.26 (−2.98)	0.00 (−5.21)	0.25	0.00	0.86	1.84	7.48	1357.56	0.33
3	80.49 (11.56)	0.0002 (8.90)	−0.30 (−4.29)	−0.02 (−6.18)	0.26	0.00	0.90	1.87	8.25	346.34	0.37
5	90.56 (17.75)	0.0005 (11.46)	−0.33 (−5.09)	−0.04 (−6.50)	0.26	0.00	0.85	1.95	7.41	207.95	0.39
7	94.17 (18.46)	0.0010 (14.32)	−0.35 (−5.24)	−0.08 (−7.20)	0.27	0.01	0.91	2.00	7.12	139.98	0.42
9	102.32 (22.06)	0.0014 (16.23)	−0.38 (−5.68)	−0.11 (−7.10)	0.28	0.01	1.22	2.05	7.38	118.57	0.41
11	105.91 (27.24)	0.0025 (18.66)	−0.43 (−6.69)	−0.17 (−6.53)	0.29	0.01	0.70	2.07	6.02	86.82	0.43
13	104.48 (23.31)	0.0033 (17.61)	−0.44 (−6.75)	−0.24 (−7.11)	0.30	0.02	1.17	2.11	6.09	68.62	0.41
15	103.18 (25.53)	0.0043 (18.93)	−0.46 (−7.33)	−0.31 (−6.89)	0.32	0.02	1.29	2.16	5.50	61.65	0.42
17	105.65 (25.54)	0.0060 (19.64)	−0.46 (−7.10)	−0.39 (−6.14)	0.36	0.03	1.23	2.15	4.48	50.37	0.42
19	105.55 (27.12)	0.0087 (18.45)	−0.48 (−7.83)	−0.62 (−6.81)	0.41	0.05	1.23	2.14	4.02	42.00	0.45
21	101.25 (22.93)	0.0120 (19.48)	−0.51 (−8.44)	−0.78 (−6.19)	0.47	0.07	1.55	2.24	3.85	34.95	0.44
23	110.67 (20.67)	0.0175 (16.60)	−0.56 (−8.14)	−1.08 (−5.39)	0.58	0.10	1.60	2.37	3.35	29.97	0.47
25	117.01 (16.72)	0.0432 (20.25)	−0.64 (−8.31)	−3.13 (−7.17)	0.99	0.24	2.05	2.69	2.80	15.86	0.47
Panel C:	1990—2006										
Portfolio	β^{1p} (·100)	β^{2p} (·100)	β^{3p} (·100)	β^{4p} (·100)	$E(C^p)$ (%)	$\sigma(\Delta c^p)$ (%)	$E(R^p - R^f)$ (%)	$\sigma(R^p)$ (%)	*TRN* (%)	*CAP* (billion ¥)	*BM*
1	58.55 (13.94)	0.0000 (5.42)	−0.47 (−4.59)	0.00 (−7.82)	0.25	0.00	0.25	2.00	5.63	2468.08	0.46
3	78.87 (25.48)	0.0003 (8.58)	−0.53 (−4.74)	−0.01 (−8.24)	0.26	0.00	0.21	2.18	5.85	508.02	0.53
5	86.23 (30.75)	0.0008 (11.87)	−0.65 (−5.68)	−0.03 (−8.23)	0.27	0.00	0.12	2.29	5.90	254.24	0.57
7	90.88 (38.66)	0.0014 (13.63)	−0.64 (−5.42)	−0.06 (−8.71)	0.28	0.01	0.12	2.32	4.99	179.04	0.66
9	98.07 (41.65)	0.0028 (16.04)	−0.69 (−5.51)	−0.11 (−8.92)	0.30	0.01	0.19	2.38	5.21	128.67	0.70
11	100.81 (44.67)	0.0047 (15.60)	−0.73 (−5.73)	−0.16 (−7.13)	0.33	0.02	0.08	2.36	5.08	91.14	0.73
13	105.38 (48.09)	0.0084 (19.60)	−0.83 (−6.38)	−0.28 (−7.72)	0.38	0.04	0.07	2.38	5.06	75.05	0.83
15	102.02 (46.66)	0.0128 (18.01)	−0.83 (−6.60)	−0.38 (−6.45)	0.44	0.06	0.22	2.39	4.75	63.68	0.86
17	100.37 (49.02)	0.0206 (21.35)	−0.79 (−6.32)	−0.54 (−6.03)	0.54	0.09	−0.02	2.35	3.83	51.01	0.92
19	104.90 (41.47)	0.0301 (24.71)	−0.80 (−6.04)	−0.85 (−6.90)	0.68	0.13	0.22	2.43	3.96	38.54	0.94
21	98.50 (41.89)	0.0528 (31.05)	−0.81 (−6.63)	−1.13 (−5.22)	0.89	0.22	0.26	2.44	3.63	36.46	0.98
23	106.48 (42.62)	0.0823 (27.34)	−0.83 (−6.18)	−1.74 (−5.01)	1.35	0.36	0.42	2.64	3.35	27.88	1.00
25	115.38 (37.44)	0.2856 (33.93)	−0.99 (−6.82)	−6.01 (−5.19)	3.32	1.20	0.58	3.03	2.34	17.04	0.96

Table 8

Correlations for portfolio betas

This table reports the correlation matrices of the four betas, β^{1p}, β^{2p}, β^{3p}, and β^{4p}, for the 25 equally weighted illiquidity-sorted portfolios.**, and * denote significance at the 1% and 5% levels, respectively.

Variable		β^{2p}	β^{3p}	β^{4p}
Panel A:	1975–2006			
β^{1p}		0.529**	−0.938**	−0.544**
β^{2p}			−0.670**	−0.997**
β^{3p}				0.685**
Panel B:	1975–1989			
β^{1p}		0.558**	−0.888**	−0.544**
β^{2p}			−0.818**	−0.998**
β^{3p}				0.799**
Panel C:	1990–2006			
β^{1p}		0.529**	−0.936**	−0.538**
β^{2p}			−0.645**	−0.996**
β^{3p}				0.657**

period. Assuming that $\lambda = E(R_t^M - C_t^M - R^f)$ is the same for all betas in the equation, we define the 'net beta' as

$$\beta^{np} = \beta^{1p} + \beta^{2p} - \beta^{3p} - \beta^{4p} \tag{14}$$

and the liquidity-adjusted CAPM reduces to

$$E\left(R_m^p - R_m^f\right) = \alpha + \kappa E\left(C_m^p\right) + \lambda\beta^{np} \tag{15}$$

As a comparison, we also estimate the standard CAPM

$$E\left(R_m^p - R_m^f\right) = \alpha + \lambda^1\beta^{1p} \tag{16}$$

The estimation results for equations (15) and (16) are shown in rows and 2 of Table 9, respectively. For equation (15), the estimated coefficient of the illiquidity level is 0.403, with a highly significant t-value of 4.943; the net beta coefficient is 0.003, with a marginally significant t-value of 1.795. For equation (16), the estimated coefficient for the market beta is 0.010, with a t-value of 2.743, which is significant at the 1% level. These results suggest that both liquidity-adjusted and standard CAPMs are supported by the data. However, the goodness of fit is much better for the liquidity-adjusted CAPM, since the adjusted R^2 is 0.82 for the estimated equation (15), but only 0.38 for equation (16). These results are again very consistent with the corresponding results reported by Acharya and Pedersen (2005) using US data.

Next, we isolate the effect of liquidity betas from the level of liquidity $E(C_m^p)$ and the market beta β^{1p} by testing

$$E\left(R_m^p - R_m^f\right) = \alpha + \kappa E\left(C_m^p\right) + \lambda^1\beta^{1p} + \lambda\beta^{np} \tag{17}$$

$$E\left(R_m^p - R_m^f\right) = \alpha + \lambda^1\beta^{1p} + \lambda\beta^{np} \tag{18}$$

Table 9

A test of the liquidity-adjusted CAPM on portfolio returns

This table reports the estimated coefficients from GMM regressions of the liquidity-adjusted CAPM for 25 equally weighted illiquidity-sorted portfolios using monthly data from July 1975 to December 2006. The liquidity-adjusted CAPM is tested by running excess monthly portfolio returns on the monthly normalized portfolio illiquidity, the market beta, and illiquidity betas:

$$E\left(R_m^p - R_m^f\right) = \alpha + \kappa E\left(C_m^p\right) + \lambda^1\beta^{1p} + \lambda^2\beta^{2p} + \lambda^3\beta^{3p} + \lambda^4\beta^{4p}$$

where $E(C_m{}^p)$ is the expected level of illiquidity for portfolio p, $\beta^{1p}$is the market beta, while β^{2p},β^{3p}, and β^{4p} are the liquidity betas for portfolio p. We assume the risk premium is equal for all betas, $\lambda = \lambda^1 = \lambda^2 = \lambda^3 = \lambda^4$, and the net beta $\beta^{np} = \beta^{1p} + \beta^{2p} - \beta^{3p} - \beta^{4p}$. The t-statistics are reported in parentheses.**, and * denote significance at the 1% and 5% levels, respectively.

	Constant	$E(C^p)$	β^{1p}	β^{2p}	β^{3p}	β^{4p}	β^{np}	Adj. R^2
Panel A: 1975–2006								
1	0.002	0.403**					0.003	0.820
	(1.788)	(4.943)					(1.795)	
2	−0.003		0.010**					0.377
	(−0.925)		(2.743)					
3	0.000	1.171**	0.273**				−0.268**	0.851
	(0.394)	(5.227)	(3.353)				(−3.316)	
4	0.003		−0.134**				0.136**	0.780
	(1.689)		(−4.084)				(4.315)	
5	0.001	1.301**	−0.006	−1.347	−1.041*	0.305**		0.895
	(1.212)	(4.980)	(−1.350)	(−0.414)	(−2.036)	(4.218)		
Panel B: 1975–1989								
1	0.000	1.594**					0.006*	0.778
	(0.196)	(4.639)					(2.287)	
2	−0.005		0.018**					0.393
	(−1.033)		(3.332)					
3	−0.010	5.696*	0.957				−0.947	0.811
	(−1.539)	(2.266)	(1.690)				(−1.679)	
4	0.004		−0.353**				0.358**	0.748
	(1.676)		(−3.899)				(4.051)	
5	0.002	1.439	0.003	121.262	−0.273	1.689		0.824
	(0.127)	(0.235)	(0.532)	(0.623)	(−0.207)	(1.107)		
Panel C: 1990–2006								
1	0.004**	0.201**					−0.003**	0.728
	(4.944)	(7.109)					(−3.419)	
2	0.000		0.002					0.056
	(−0.108)		(0.917)					
3	0.003**	0.647**	0.225**				−0.227**	0.802
	(4.654)	(6.962)	(5.306)				(−5.344)	
4	0.004**		−0.094**				0.091**	0.650
	(3.992)		(−6.083)				(6.120)	
5	0.003**	0.837**	0.000	−3.728	0.464	0.150*		0.810
	(3.732)	(5.263)	(−0.198)	(−1.277)	(1.665)	(2.128)		

The estimation results are presented in rows 3 and 4 of Table 9, respectively. While the estimated coefficients for illiquidity level and the market beta are still positive and significant, the estimated coefficient for net beta is negative and significant in equation (17). This suggests that liquidity risk matters after controlling for the illiquidity level and market risk. However, a negative estimated coefficient for β^{np} in equation (17) is counterintuitive, since the risk premium should not be negative. For equation (18), the estimated coefficient is negative for β^{1p} but positive for β^{np}, and both are significant. Acharya and Pedersen (2005) show that a negative coefficient for β^{1p}does not imply a negative risk premium on market risk, since $\beta^{1p}$is also contained inβ^{np}. Instead, a negative coefficient suggests that liquidity risk may have a higher risk premium than market risk. Overall, our results support, although not very strongly, the view that liquidity risk matters more than the effect of the illiquidity level and the market risk.

Finally, we estimate an unconstrained version that allows different risk premiums for different betas:

$$E\left(R_m^p - R_m^f\right) = \alpha + \kappa E\left(C_m^p\right) + \lambda^1\beta^{1p} + \lambda^2\beta^{2p} + \lambda^3\beta^{3p} + \lambda^4\beta^{4p} \qquad (19)$$

In row 5 of panel A of Table 9, we see that the estimated κ is significantly positive, indicating the illiquidity level affects stock returns. The estimated risk premiums for β^{1p}andβ^{2p} are insignificant and negative. While the estimated risk premium for $\beta^{3p}$is negative and significant as expected, it is positive and significant forβ^{4p}, which is counterintuitive. However, no strong inference can be made due to severe multicollinearity among the four betas.

Since the liquidity effect on stock returns differs across subsample periods in the previous sections, we repeat the GMM regressions for the two subsample periods 1975–1989 and 1990–2006 and report the results in panels B and C of Table 9, respectively. They are qualitatively the same as those reported in Panel A of Table 9.

6. Summary and Conclusion

With some modification of the illiquidity measure proposed by Amihud (2002), we conduct a study on the relation between illiquidity, illiquidity risk, and stock returns on the TSE. We first examine the cross-sectional relation between illiquidity and stock returns and find that illiquidity has a positive impact on stock returns in Japan in general, but not in the second subsample period of 1990–2006. Even after deleting the monthly observations associated with negative market premiums, we still fail to find a significant relation between illiquidity and stock returns in the second subsample period. We conjecture that some institutional and behavioural changes may have happened during that period. Further studies are needed to explore this issue.

Next, we look at the time series relation between illiquidity and stock returns. While the expected illiquidity has a positive and significant impact on expected stock returns, the unexpected illiquidity has a negative and significant impact on contemporaneous stock returns. In addition, we find some evidence in support of the ‘flight to liquidity’ hypothesis. Separating the whole sample into two subperiods produces similar results. These findings are largely consistent with those found by Amihud (2002) using US data.

Finally, we use the liquidity-adjusted CAPM framework proposed by Acharya and Pedersen (2005) to examine if liquidity risk is priced over and above the liquidity level and the market risk. We find that the liquidity-adjusted CAPM is superior to the standard

CAPM. However, we only obtain weak evidence for the argument that liquidity risk is priced in addition to the liquidity level and the market risk.

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