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Citation:

Chowdhury, A and Uddin, M and Anderson, K (2017) Liquidity and macroeconomic management in emerging markets. Emerging Markets Review. ISSN 1566-0141 DOI: <https://doi.org/10.1016/j.ememar.2017.10.001>

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Liquidity and Macroeconomic Management in Emerging Markets

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Abstract

Emerging markets have received considerable attention for foreign investment and international diversification due to the possibility of higher earnings and a low level of integration with global equity markets. These high returns often need to be balanced by the high liquidity costs of trading in illiquid emerging markets. Several studies have shown that central bank and government policies are significant determinants of market liquidity. We investigate the influence of monetary and fiscal policy variables on the market and firm level liquidity of eight emerging stock markets of Asia. Using four different (il)liquidity measures and nine macroeconomic variables, we find that changes in the money supply, government expenditure and private borrowing significantly affect stock market liquidity. Illiquidity is also strongly affected by the bank rate, short-term interest rate and government borrowing. We demonstrate that ‘crowding out’ and ‘cost of funds’ effects exist in these markets. Other major findings are that some markets are more sensitive to local macroeconomic news than world factors, the impact on size based portfolios largely depends on the instruments used by the central banks and government, the liquidity of the manufacturing sector is affected by changes in any policy variables, financial institutions are only influenced by monetary policy variables, and the service sector is least affected.

Keywords: Liquidity, Macroeconomics, Emerging markets, Central bank, SE Asia, crowding out

JEL classification: G15, G18, G28

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This paper was presented at the 2016 Cross Country Perspectives of Finance conferences held in Taiyuan and Pu'er, China.

1. Introduction

The higher returns often available from investing on emerging markets need to be balanced by their liquidity costs. Governments' monetary and fiscal policies play a central role in determining this liquidity. While the effects of macroeconomic policies on stock market liquidity have been thoroughly explored, there has been little or no work on how the policies of emerging market governments affect liquidity. In this paper we look at eight emerging markets in SE Asia and show that several macroeconomic policies affect liquidity, but that causality can operate in the other direction too (a successful stock market is often seen in these markets as an indicator of a successful ruling party). Some markets behave a lot like developed markets in their sensitivity to world news but others, particularly Bangladesh and Pakistan, march to their own tune. We also show that different industry sectors are affected by different policy variables.

The role of liquidity in stock markets and economic development has been documented in a number of recent pieces of research. For example, illiquidity shocks could be a reason for recessions and stock market crashes (Jaccard, 2013). Liquidity has been identified as a leading indicator of real economy (Næs et al., 2011) and is thought to be a reliable predictor of future economic growth (Levine and Zervos, 1989). Stock market liquidity may also be used as a proxy for investors' liquidity and transaction costs (Goyenko et al., 2009). Considering these important interconnections between stock market liquidity and both macro and micro market structure, Choi and Cook (2006) argue that the unpredictability of market liquidity is an important source of risk for equity investors. Due to the immense importance of stock market liquidity, both government and market regulators are making constant efforts to maintain a satisfactory level of liquidity. One of the major ways to influence market liquidity is via macroeconomic policy tools, both monetary and fiscal. A number of studies such as Fernández-Amador et al. (2013), Brunnermeier and Pedersen (2009), Easley and O'Hara (1987), Birz and Lott (2011), Chordia, Roll and Subrahmanyam (2001), Goyenko and Ukhov (2009) and Choi and Cook (2006) have explained the influence of economic policies on stock market liquidity. However, one notable similarity in all those studies is that they have examined market liquidity using developed stock markets. This is against the backdrop that emerging markets are gaining importance and they are notably different from developed markets in terms of their institutional and regulatory set ups (Bekaert et al., 2007). Due to this difference between developed and emerging markets, results from developed markets may not be generalised for emerging markets. Bekaert et al. (2007) further mention that the focus on emerging markets should yield powerful tests and useful independent evidence. In addition, the relationship of market liquidity with fiscal policy is still

significantly under-explored and ambiguous with only the one exception of Gagnon and Gimet (2013) who examined the effect of fiscal policy on market liquidity using US macroeconomic data.

In this study, therefore, our objective is to investigate the influence of monetary and fiscal policy variables on the liquidity of eight emerging stock markets. We examine whether the monetary policies of central banks and the fiscal policies of governments are common determinants of stock liquidity. For example, when the central bank pursues an expansionary monetary policy, the increase in funds could cause higher order inflows into the stock market and potentially change liquidity (see Choi and Cook, 2006; Chordia et al., 2005). Moreover, due to any systematic risk or information shock (e.g. macroeconomic policy uncertainty) investors might change their asset holdings between stocks and other financial securities. In the first step, therefore, we observe the impact of standard monetary and fiscal policies on aggregate stock market liquidity. In the second step, we extend our analysis to look more deeply into the influence of those policies on the liquidity of sectors and individual stocks.

As our second objective we examine whether the effect of macroeconomic policy depends on the size of firms. Amihud (2002) finds that small stocks are more responsive to illiquidity shocks and large firms become more attractive when aggregate liquidity declines. In contrast, Fernández-Amador et al. (2013) report smaller firms are more responsive to liquidity shocks and the liquidity-providing effect of e.g. a loose monetary policy is stronger for larger firms. Similarly, Næs et al. (2011) assert that the informativeness of stock market liquidity is highest in smaller firms, which are the least liquid. In this study we further investigate this linkage for firms listed on emerging stock markets.

Finally, we investigate the characteristics of liquidity in emerging markets during the 2007-08 financial crisis. Many studies have reported that market liquidity dropped significantly during the crisis (see e.g. Blanchard et al., 2010; Söderberg, 2008; Choi and Cook, 2006). We therefore want to see how policies adopted by governments during the financial crisis influenced the flow of funds in their financial markets.

Our contributions are threefold. First, by considering macroeconomic management variables as determinants and examining their influence, this paper helps to identify the commonality observed or underlying economic forces responsible for the systematic movement of liquidity in emerging markets.

Second, we know very little about the dynamic relationship of monetary and fiscal policy with stock market liquidity. Yet studies such as Spilimbergo et al. (2009), Blanchard et al. (2010) and Woodford (2011) have detailed how fiscal stimulus packages inject liquidity into financial markets. We will therefore provide new insights into the interaction of these variables in emerging economies. As Choi and Cook (2006) assert, the unpredictability of market liquidity is an important source of risk for investors, but it is clear that the risk is higher for emerging markets where investors generally have less opportunity to diversify their portfolios and face greater asymmetry of information. Consequently, identifying the macroeconomic determinants of the liquidity of these markets will help both local and international investors.

Third, the reputation of emerging stock markets in the context of investment portfolios and international diversification has received considerable attention over the last decade when developed markets have faced financial crisis. In particular, emerging markets are experiencing explosive growth fuelled by foreign investment and offered returns up to 90 percent (Lesmond, 2005). The United Nations Conference on Trade and Development reports that developing economies attracted 795 billion US\$ of foreign investment in 2013, which is 52 percent of total foreign direct investment in the world (UNCTAD, 2014). Ten emerging markets are now ranked among the top 20 markets in the world and account for more than 30 percent of world market capitalization. According to the International Monetary Fund the growth rate of emerging markets was slowing in 2015 but is expected to pick up in 2016 (WEO, 2015). Hence researchers, investors and policy makers would be interested to know the characteristics of market liquidity in these markets, specifically during the 2007-8 financial crisis.

Emerging markets have different economic and institutional characteristics from developed markets that make them interesting for research as highlighted in e.g. Batten and Szilagyi (2011). They are largely dominated by individual investors, have low integration with world markets and a significant degree of political uncertainty. In many of these economies, stock market performance is seen as an indicator of political success. Therefore, governments often intervene in these markets to ensure future electoral success. These markets generally have small scale bond markets but strong alternative investment opportunities (e.g. national savings deposits and wage earners' bonds). These distinctive characteristics of emerging markets make it worth investigating the dynamic relationship between market liquidity and the policies of emerging market central banks and governments.

Our major findings include: expansionary (restrictive) monetary and fiscal policy has a positive impact on equity market liquidity (illiquidity). There is causality and bidirectional causality

between macroeconomic variables and (il)liquidity ratios. Bank rate, short-term interest rate and government borrowing can explain a large fraction of the error variance of Amihud's (2002) illiquidity measures and turnover price impact ratio. This indicates the existence of 'crowding out' and 'cost of funds' effects in these stock markets. The global financial crisis broadly had a negative influence on market liquidity and significantly increased market illiquidity. Individual stock characteristics such as market capitalization play an important role in the relationship between firm level (il)liquidity and central bank or government policies, such as smaller and larger firms being more sensitive to changes in any variables. For sectoral portfolios, financial institutions in this region are strongly associated with central bank policies, whereas the manufacturing sector is subject to both. The liquidity response to macroeconomic shocks in Taiwan is the most consistent among the eight equity markets and well in line with theoretical arguments. On the other hand, Pakistan does not show any clear trend.

The rest of the study is organised as follows: section 2 reviews the methodological framework and details of the data. Empirical results are reported in section 3. Section 4 summarizes the findings and makes concluding remarks.

2. Data and Methodology

2.1 Data set

Our sample consists of financial and macroeconomic data of eight Asian emerging countries (i.e. Bangladesh, India, Indonesia, Malaysia, Pakistan, South Korea, Taiwan and Thailand) for a sixteen year window starting from January 2000 and ending in December 2015, which is 182 months each for about 1050 selected firms. This is equal to samples of one million daily and 191,100 monthly observations. In addition, there are about 10,000 monthly macroeconomic data items from these eight markets. The main source of this information is Thomson Reuters Datastream. To compute the returns, liquidity and illiquidity measures stocks are included or excluded based on the criteria stated in Chordia et al. (2005) and Fernández-Amador et al. (2013). First, to be included, a stock has to be present at the beginning and at the end of the year; second, it must have traded more than 100 trading days during the calendar year; third, a stock has to be listed for at least fifteen years on the exchange; fourth, to avoid the influence of outliers, we exclude the sample company if the price at any month-end during the year is greater than 1000 in the local currency; finally, due to their different trading characteristics from ordinary shares, we also exclude assets in the following categories – commercial bonds, treasury bonds, mutual funds (both open and closed end), and preferred stocks.

2.2 Liquidity and illiquidity measures

Liquidity is an elusive concept and is not observed directly but rather has a number of aspects that cannot be captured in a single measure (Amihud, 2002). Following the procedures suggested by Goyenko and Ukhov (2009), Amihud (2002) and Datar, Naik and Radcliffe (1998), this study uses four different measures to capture the aspects of trading activity and price impact.

The first proxy of liquidity for an asset that we use in this study is turnover rate (TR), as suggested in Datar et al. (1998). The turnover rate of a stock is the number of shares traded divided by the number of shares outstanding in the stock. This is an intuitive metric of the liquidity of the stock. The relationship between trading volume and market liquidity is highlighted in much previous literature, such as Amihud and Mendelson (1980), and Lo and Wang (2000). Fernández-Amador et al. (2013) assert that stock turnover can be interpreted as the reciprocal of the average holding period, which means stocks with higher turnover are on average held for shorter time periods, and thus exhibit increased trading activity. The second variable we use as a proxy of liquidity is traded volume (TV). The relationship is documented in Brennan et al. (1998) – higher trade volume implies increased liquidity. They find a robust effect of trading volume with the presence of both risk-adjusted and unadjusted returns, which supports the notion that this variable is acting as a proxy for the liquidity of the market in the firm's share (see Brennan et al., 1998 for a detailed discussion).

The mathematical measure for each category (i.e. TR and TV) of liquidity proxies can be expressed as follows:

$$TR_{iym} = \frac{\sum_{d=1}^{D_{iym}} VO_{iymd}}{NSO_{iym}} \quad (i)$$

where TR_{iym} is the turnover rate of stock i in month m of year y ; $\sum_{d=1}^{D_{iym}} VO_{iymd}$ is the monthly sum of the daily number of shares traded and NSO_{iym} is the number of shares outstanding; and

$$TV_{iym} = \ln \left(\sum_{d=1}^{D_{iym}} (VO_{iymd} P_{iymd}) \right) \quad (ii)$$

where TV_{iym} is the traded volume of stock i in month m of year y ; VO_{iymd} is the number of daily traded shares and P_{iymd} is the daily price of each share. Therefore, traded volume is calculated by taking the natural logarithm of the monthly sum of the daily product of the number of shares traded and their respective market price. Both TR and TV are based on trading activity

and we can interpret them as liquidity proxies, as higher values are associated with more liquid assets.

Our third and fourth measures are of illiquidity rather than liquidity. Our third measure is Amihud's (2002) illiquidity ratio, which quantifies the response of returns to one dollar of trading volume. This illiquidity measure is very well established, particularly since studies such as Hasbrouck (2009) and Goyenko et al. (2009) have reported its efficacy as a measure of price impact. For our fourth measure we use turnover price impact ratio (TPI), which was proposed and developed by Florackis et al. (2011). The ratio can be defined as the returns impact of a one percent stock turnover. It has several appealing characteristics compared to Amihud's (2002) illiquidity ratio. TPI is free from size bias and considers trading frequency rather than volume, thus it is less related to market capitalization and can encapsulate stocks' cross-sectional variability (see Florackis et al., 2011 for a detailed discussion). Fernández-Amador et al. (2013) suggest that TPI is more isolated from nominal effect than Amihud's (2002) ILLIQ, therefore it can offer different conclusions in an environment where nominal shocks dominate (e.g. an inflationary environment).

The mathematical expressions for ILLIQ and TPI are as follows:

$$ILLIQ_{iyd} = \frac{|R_{iyd}|}{TV_{iyd}} \quad (iii)$$

where $ILLIQ_{iyd}$ is the illiquidity ratio of security i on day d of year y ; R_{iyd} is the return on stock i on day d of year y and TV_{iyd} is the respective daily volume; and

$$TPI_{iyd} = \frac{|R_{iyd}|}{TR_{iyd}} \quad (iv)$$

where TPI_{iyd} is the turnover price impact of security i on day d of year y ; TR_{iyd} and R_{iyd} is the turnover rate and daily return of each share respectively. For our empirical models we compute the monthly average of the individual liquidity or illiquidity of each stock and also the equally weighted cross-sectional averages of the liquidity and illiquidity for the vector autoregression model.

2.3 Macroeconomic variables

The prime objective of this study is to investigate the effect of monetary policy (MP) and fiscal policy (FP) on the liquidity of selected emerging stock markets. To achieve this objective we select several monetary and fiscal policy variables in line with previous studies, e.g. Chordia, Roll

and Subrahmanyam (2001), Chordia et al. (2005), Goyenko and Ukhov (2009). We approximate the monetary policy by using data on the aggregate money supply, bank rate and short-term interest rate. For aggregate money supply we use the rolling twelve month growth rate of base money (BM). Our second variable, bank rate (BR), is relatively straightforward and is the rate at which the central bank offers credit to other FIs and thus acted as a control mechanism for market money supply. Finally, we take the short-term interest rate (SIR) to consider domestic interest rate shocks on liquidity following the argument of Chordia, Roll and Subrahmanyam (2001) and Söderberg (2008). The three-month Treasury bill rate is used as the proxy for this short-term interest rate, as suggested in Gagnon and Gimet (2013).

Next to monetary policy, we consider three fiscal policy variables to capture the government's intervention in equity market liquidity. These variables are government expenditure (GE), government borrowing from commercial banks (GB) and credit to the private sector (CP). The recent literature identifies a chain of transmission of budgetary policies to the credit sector (Gagnon and Gimet, 2013). For example, Blanchard (2009) reports that the financial crisis affected the economy through credit rationing, i.e. tightening of lending standards by banks. Credit to the private sector therefore measures the bank's lending volume and willingness to loosen the credit standard (e.g. margin requirements) following a fiscal policy shock (see Gagnon and Gimet, 2013). For our first fiscal variable, Blinder and Solow (1973) assert that a dollar of additional government spending raises national income not only by the original dollar but has a multiplier effect of perhaps several dollars in the economy. This can facilitate the money flow in the economy and thus any positive shock may increase liquidity. On the other hand, Fisher (1988) among many others states that borrowing from commercial banks by the government can create a 'crowding out' effect and thus create competition for private savings where business firms may suffer from lack of credit opportunities.

Other macroeconomic factors, such as unexpected productivity falls and excessive inflationary pressures, are likely to influence illiquidity indirectly by inducing fund outflows, reducing price, increasing volatility and exacerbating inventory risk (Goyenko and Ukhov, 2009). The interrelationship between various macroeconomic variables (i.e. other than monetary and fiscal policy variables) and market liquidity is theoretically developed in Eisfeldt (2004) and also empirically studied and documented in Söderberg (2008), Næs et al. (2011) and Fernández-Amador et al. (2013). Based on their procedure and suggestions, we include monthly growth rate of industrial production (IP) and monthly inflation rate (IR) to capture business cycle and inflation development. Further to control individual stock characteristics, following the argument

of Fernández-Amador et al. (2013), Brunnermeier and Pedersen (2009) and Copeland and Galai (1983), we include monthly return (RET) and standard deviation (STD). We compute both return and standard deviation from equally weighted averages of individual monthly stock prices. Finally, to control for the size effect of firms and cyclical movements of the equity market we consider the log value of market capitalization (lnMV) and the benchmark stock index (IDX) respectively in our empirical model (see Fernández-Amador et al., 2013 for more discussion).

2.4 Empirical Models

2.4.1 Vector Autoregression Analysis

Our aim is to understand the relationship between equity market liquidity and its primitive drivers, macroeconomic management variables. Based on earlier studies, we expect possible endogenous associations between market liquidity, monetary, fiscal and other macroeconomic policies. However, there is good reason to assume bidirectional causality as well among various market characteristics and liquidity, as suggested by Chordia et al. 2005; Fernández-Amador et al., 2013; Gagnon and Gimet, 2013).

Given the bidirectional causalities, we investigate the association between macroeconomic management variables and market liquidity using the vector autoregression procedure employed in Chordia et al. (2005) and Goyenko and Ukhov (2009). Here, the following system is considered:

$$X_t = c + \sum_{j=1}^k B_{1j} X_{t-j} + u_t \quad (v)$$

where X_t is a vector that represents endogenous variables - liquidity, returns, volatility, industrial production, inflation, monetary and fiscal policy instruments; c is the vector of intercept, B is a six \times six coefficients matrix (for the monetary and fiscal policy variables), and u_t labels the vector of residuals. The number of lags is estimated based on Akaike information criterion and the Schwarz information criterion. Where they indicate different lag lengths, we choose the lesser lag length for parsimony (see Chordia et al. 2005). The augmented Dickey-Fuller (1979) test is used to check the non-stationarity of the variables.

Finally, within this VAR system, we check the influence of the financial crisis (2007-08) on each liquidity measure by adding a dummy variable for the crisis. To get a clear and comparable picture of the impact of the crisis, we split the sample between the crisis 2007-08 and non-crisis periods. These results will therefore indicate which Asian equity markets were affected by the

subprime mortgage crisis 2007-08 originating in the US. It will also help policy makers of the respective country to formulate future strategies around their market sensitivity.

2.4.2 Panel regression for individual stocks

Having assessed the impact on overall market liquidity, in our second step we look into the influence of monetary and fiscal policy stance on liquidity of individual firms and sectoral stocks. For this purpose, we combine time-series and cross-sectional data of liquidity, returns, volatility, industrial production, inflation, monetary and fiscal policy instruments to estimate panel regressions as suggested in Fernández-Amador et al. (2013). The liquidity ($LIQ_{i,t}$) of stock i in month t is modelled as a function of the one-month lagged macroeconomic management variables and other lagged control variables. We model separately for monetary (i.e. equation vi) and fiscal policy (i.e. equation vii). The models are:

$$LIQ_{i,t} = c + \alpha_1 LIQ_{i,t-1} + \alpha_2 MP_{t-1} + \alpha_3 Int_{i,t-1} + b_4 Y_{i,t-1} + \alpha_5 Z_{i,t-1} + c_i + u_{i,t} \quad (vi)$$

$$LIQ_{i,t} = c + \alpha_1 LIQ_{i,t-1} + \alpha_2 FP_{t-1} + \alpha_3 Int_{i,t-1} + b_4 Y_{i,t-1} + \alpha_5 Z_{i,t-1} + c_i + u_{i,t} \quad (vii)$$

where $LIQ_{i,t}$ is the dependent variable and represents the four (il)liquidity ratios. To account for the possible autocorrelation induced by a dynamic relationship in stock liquidity, we include one-month lagged (il)liquidity measures $LIQ_{i,t-1}$ as regressors. MP_{t-1} and FP_{t-1} are monetary policy and fiscal policy variables respectively. We consider an interaction term $Int_{i,t-1}$ in both equations that indicates whether the influence of monetary policy or fiscal policy depends on size of firm as measured by log market capitalization, $lnMV_{i,t-1}$. Therefore, the interaction term is $(MP * lnMV_{i,t-1})$ in equation (vi) and $(FP * lnMV_{i,t-1})$ in equation (vii). However, we do not use any such interaction term in our models for sectoral stocks. The vector $Y_{i,t-1}$ denotes lag values of our other control variables for stock characteristics; they are monthly returns ($RET_{i,t-1}$), monthly standard deviation of daily stock returns ($STD_{i,t-1}$) and natural logarithm of market capitalization ($lnMV_{i,t-1}$). The macroeconomic variables to capture cyclical variation, such as monthly growth rate of industrial production ($IP_{i,t-1}$), inflation rate ($IR_{i,t-1}$) and share price index ($IDX_{i,t-1}$) are represented by the vector $Z_{i,t-1}$. As recommended in Fernández-Amador et al. (2013), to account for time-invariant stock-specific determinants of liquidity, we use the fixed-effect (within) estimator and c_i is the fixed-effect in this cross-section. Finally, $u_{i,t}$ is the residual in our models. The number of lag order for each variable is selected based on whether any autocorrelation exists in the residuals.

3. Empirical results

The empirical results on the association of market liquidity with monetary and fiscal policy variables are reported in this section. Section 3.1 presents the preliminary statistics of our cross-sectional and time series data, bivariate cross-sectional correlation among mean of time series data and panel cointegration between macroeconomic management variables and stock price. Section 3.2 discusses the influence of monetary policy, fiscal policy and the 2007-8 financial crisis on aggregate market liquidity. Finally, firm and industry level evidence on the connection between liquidity and macroeconomic variables is assessed in section 3.3.

3.1 Preliminary statistics

We present the summary statistics in Tables 1 and 2, where Table 1 shows the statistics for cross-sectional variables that we use in our panel estimations and Table 2 includes descriptive statistics related to time-series macroeconomic variables. However, our particular interest is to see the bivariate correlation between VAR endogenous variables and the average monthly values of the four (il)liquidity ratios. The results are given in Table 3. It is obvious that one could expect cross-sectional correlation between the trading activity measures (i.e. TR and TV) and the price impact ratios related to illiquidity (i.e. ILLIQ and TPI). Fernández-Amador et al. (2013) assert that this observation is intuitive, since higher trading activity translates into more liquid stocks, whereas higher levels of price impact or transaction costs indicate less liquid stocks. From the median values of the correlation matrix of each market, we find positive correlation respectively between trading activity and price impact ratios, and yet their cross-sectional relationships are found to be negative as expected. For example, among these four (il)liquidity measures, the highest negative correlation is reported between trading volume (TV) and Amihud's illiquidity ratio, which is -0.112. Similarly, the highest positive correlation is identified between Amihud's (2002) illiquidity ratio (i.e. ILLIQ) with the turnover price impact (0.117). These results are also statistically significant at the 1 percent level.

Other than the bidirectional correlation between liquidity measures, Table 3 also reports their relations with return (RET), volatility (STD) and market capitalization (lnMV). The positive (negative) correlation between the market value of firms and liquidity (illiquidity) suggests that stocks of larger firms tend to be more liquid as suggested in Fernández-Amador et al. (2013). Further from the median values in Table 3, we find that TR (0.13) and TV (0.04) are positively related to lnMV but there is a negative association with ILLIQ (-0.05) and TPI (-0.05).

Table 1: Descriptive statistics for cross-sectional variables

	Mean of monthly mean	Median of monthly mean	Max. of monthly mean	Mini. of monthly mean	Mean of monthly σ	Mean of monthly skewness
TR	2.3259	2.1497	5.1484	1.1914	0.7966	1.8144
TV	12.6894	12.9613	15.1625	8.7732	1.6242	-0.8117
ILLIQ	0.0014	0.0005	0.0351	3.09E-05	0.0036	6.1442
TPI	5.4719	1.1498	12.7002	0.0424	9.3148	6.1994
RET	0.0071	0.0078	0.2133	-0.2100	0.0602	-0.2964
STD	0.0884	0.0828	0.2117	0.0175	0.0308	1.0830
lnMV	11.2445	11.4553	13.7206	6.4254	1.4406	-0.9948

Note: The table represents the average of descriptive statistics for cross-sectional variables of all markets in our sample. Each variable is calculated for each stock in each month across stocks admitted to the sample in that year, and then the mean, standard deviation and skewness are calculated across stocks in each year. The table represents the mean over the sixteen years of the monthly mean, standard deviations and skewness and the median of the monthly mean, as well as the maximum and minimum monthly mean. Statistics for monthly average market capitalizations (MV) of each stock are calculated based on its natural log values.

Table 2: Descriptive statistics for macroeconomic time series

	IP	IR	lnIDX
Mean	0.5284	1.8780	7.6629
Median	0.4041	1.7086	7.7861
Maximum	18.1443	5.9750	8.5343
Minimum	-18.0782	-0.9958	6.5042
Std. dev.	4.9124	1.2363	0.5797
Skewness	-0.3511	0.5603	-0.3878

Note: The table presents the average of descriptive statistics for the macroeconomic time series of all markets in our sample. Stock index statistics are based on log values.

Table 3: Correlation matrix of time-series means of the monthly bivariate cross-sectional variables

	TR	TV	ILLIQ	TPI	RET	STD	lnMV
TR	1						
TV	0.0943*** (5.3296)	1					
ILLIQ	-0.0804*** (-6.3267)	-0.1121** (-2.0356)	1				
TPI	-0.0929*** (-7.6906)	-0.0794*** (-4.2364)	0.1178*** (13.5723)	1			
RET	0.0003 (-0.2278)	0.0417 (0.4338)	-0.0274 (-0.4082)	0.0025 (0.3457)	1		
STD	0.1355** (2.1033)	0.1110 (1.3613)	0.1649** (1.9912)	0.1455** (1.9421)	0.0878*** (-10.9269)	1	
lnMV	0.1287*** (6.2767)	0.0439*** (3.2013)	-0.0483** (-2.2938)	-0.0500 (-1.4301)	-0.0348 (-0.6386)	-0.0807*** (-2.5934)	1

Note: This table represents the median values of correlation matrix of time-series means of the monthly bivariate variables of each market.

***, ** and * are significant at the 1, 5 and 10 percent levels; t-statistics are in parenthesis.

Similar to the previous studies, such as Chordia et al. (2005) and Goyenko and Ukhov (2009) the median correlation of standard deviation with each of the liquidity measures of our sample markets is strong and statistically significant. All results are positive and the strongest association is found with the illiquidity ratio (0.32). In addition, there is a positive correlation (0.13) between return and standard deviation among the sample companies over the sample periods for all eight Asian equity markets. This suggests investors of this market could earn higher returns from risky investments. Surprisingly, the magnitude of correlations between (il)liquidity measures are relatively small, e.g. TV and TPI or TR and ILLIQ. This highlights the fact that the several (il)liquidity measures used in the analysis are not representing the same information, but different aspects of the broad concept of market liquidity (see Fernández-Amador et al., 2013).

We also check whether any long-run associations exist between the price of each stock and macroeconomic management policies applying the ‘cross-sectional cointegration’ framework of Pedroni (1999, 2004) and Kao (1999). We do this because when time series are cointegrated there must be at least one Granger causal flow in the system; moreover, the causal flow may exist because they have some other common feature (see Alexander, 2008). That means when some variables are cointegrated with each other, they may also influence some further dynamics within those series. For example, when money supply and price series are cointegrated then changes in the money supply (growth rate) may influence stock bid-ask spreads, returns or volatility. The detailed results of our panel cointegration are not reported here but available on request.

Under the structure of Pedroni (1999, 2004), we checked the long-run association of stock price respectively with monetary and fiscal policy variables. Results for each market indicate that several within-dimension and between-dimension statistics are significant from the 1 to 10 percent level. That means we can reject the null hypothesis of no cointegration and therefore, long-run associations exist between the monetary and fiscal policy variables and the monthly price of each stock in our sample. We combine the macroeconomic variables and macroeconomic management variables to run Kao’s (1999) panel cointegration with the monthly price series. That means we include the industrial production and inflation rate along with monetary and fiscal policy variables in our model. The corresponding probability of test statistics indicates that we can accept the alternative hypothesis of integration among the price series and exogenous variables at a 95 percent level of confidence.

Based on these two models, we can say that the monthly stock prices of our sample firms and macroeconomic variables of the respective country are tied together in the long term, which means in the short term they can drift apart, but over a period of time they must drift back together (see Alexander, 2008). Thus, one can emphasise that there might be causal flows between these series and their underlying characteristics, such as market liquidity and the growth rate of each macro variable may be interlinked. In the following two sections, we check that linkage using the VAR system and the panel regression model.

3.2 Influence on aggregate market liquidity

This section reports results related to the influence of macroeconomic variables on the overall liquidity of eight equity markets using the vector autoregression (VAR) framework of equation (v). We use Granger causality, impulse response function and variance decomposition associated with monetary policy, fiscal policy and the financial crisis (i.e. 2007-08) shocks. We use the

Augmented Dickey-Fuller (1979) test to check for non-stationarity. To ensure that the variables are in order of integration we employ the first difference of the variables. Each of the VAR model is estimated with the optimum lags according to the Bayesian Schwartz criterion.

3.2.1 The impact of monetary policy shocks

We estimate a total of 12 different VAR models for each stock market for each of the four (i)liquidity measures and three monetary policy variables considered in our analysis. In order to interpret the results of the estimated VAR models we also run the Granger causality test (see Granger, 1969) as suggested in Chordia (2005) and Goyenko and Ukov (2009). For the pairwise Granger causality we consider two series a_t and b_t , then estimate the equations:

$$\Delta a_t = \beta_0 + \sum_{i=1}^n \beta_{1i} \Delta a_{t-1} + \sum_{i=1}^m \beta_{2i} \Delta b_{t-1} + \varepsilon_{1t} ; \text{ and} \quad (\text{viii})$$

$$\Delta b_t = \delta_0 + \sum_{i=1}^n \delta_{1i} \Delta b_{t-1} + \sum_{i=1}^m \delta_{2i} \Delta a_{t-1} + \varepsilon_{1t} \quad (\text{ix})$$

We use an F-test for the joint significance of the coefficient assuming a null hypothesis that a_t does not Granger cause b_t and vice versa. In the null hypothesis we test that the lagged endogenous variables of interest (i.e. monetary policy or liquidity measures) do not Granger cause the dependent variable of interest (again, either market liquidity or monetary policy variables). A rejection of the null hypothesis shows the presence of Granger causality. Detailed results are not presented here but available on request.

The results indicate that there is causality and reverse causality between the monetary policy variables and the (i)liquidity ratios. Interestingly, the (i)liquidity of Indonesia, Pakistan, Taiwan and Thailand are found to be more sensitive to monetary policy. Among the monetary variables, as expected, base money growth and short-term interest rate significantly Granger cause both liquidity and illiquidity of each Asian market. In particular, base money growth Granger cause trading volume and turnover rate of Bangladesh, Indonesia, Taiwan and Thailand. Besides, reverse causality is documented in Indonesia and Thailand. For India and Pakistan base money growth rate has causality and reverse causality with both illiquidity measures - Amihud's (2002) illiquidity and turnover price impact ratio. Similarly, the short-term interest rate is found to have statistically significant causality with the illiquidity measures of Indonesia, Malaysia, Pakistan and South Korea. However, strong reverse causality exists with Amihud's (2002) illiquidity and the turnover price impact ratio of India and Thailand. Our results also identify a substantial connection of the short-term interest rate with the turnover rate (in Pakistan and South Korea) and trading volume (in India and Taiwan). Surprisingly, no causation is found in the Bangladesh

stock market for this monetary policy instrument. Finally, bank rate has causal and bidirectional causal influence only on (il)liquidity in Indonesia, Pakistan, Taiwan and Thailand. There is no pivotal impact identified for bank rate in India, Malaysia and South Korea. Overall, our findings support the cointegration reported in the previous section and favour the hypothesis that the central bank monetary policy causes the aggregate market (il)liquidity or alternatively that market (il)liquidity causes central bank monetary policy.

The Granger causality results also identify the interaction of liquidity measures with other endogenous variables. We find that stock volatility has significant causal influence on (il)liquidity in South Korea, Taiwan and Thailand. However, returns have a larger effect in Malaysia. Surprisingly, the (il)liquidity of the Indian and Pakistani equity markets has a strong impact on both returns and volatility. In particular, there is only a one-way causation between stock returns and (il)liquidity in India and these results are significant at a 99 percent level of confidence. Results indicate a strong bidirectional causality of equity returns with Amihud's (2002) illiquidity measure and turnover price impact in Pakistan and Malaysia. For the stock volatility, we see it has a two-way causation with both illiquidity ratio in India, Pakistan, South Korea and Thailand. But there is only a one-way causality between stock volatility and all four (il)liquidity measures in Taiwan. Our results, altogether, are therefore consistent with those of Chordia et al. (2005) and Goyenko and Ukhov (2009) that stock returns and volatility concurrently Granger cause market liquidity and illiquidity.

Finally, other than the (il)liquidity variables, monetary policy has a causal association with stock returns and volatility. For example, the bank rate and short-term interest rate significantly Granger cause stock returns in Indonesia, Pakistan and Taiwan. These two monetary variables Granger cause volatility in Malaysia, South Korea and Taiwan. However, in the reverse direction stock returns Granger cause the broad money growth rate in India and Indonesia; bank rate in Pakistan and Taiwan; and short-term interest rate in Indonesia. On the other hand, volatility Granger causes the bank rate in Malaysia, South Korea and Thailand; and the short-term interest rate in Indonesia, South Korea, Taiwan and Thailand.

For a deeper understanding of the dynamics of liquidity and its interaction with monetary policy within the VAR system we also report the impulse response functions (IRFs) and variance decomposition as suggested in earlier studies, e.g. Chordia et al. (2005), Goyenko and Ukhov (2009), and Gagnon and Gimet (2013). The IRF traces the impact of a one-time, unit standard deviation, positive shock to one variable on the current and future values of the endogenous variables. Results from the IRFs and variance decompositions are generally sensitive to the

specific ordering of the endogenous variables (see Chordia et al., 2005 for a detailed discussion). Therefore, in choosing an ordering, we rely on the prior evidence of Chordia et al. (2005), Goyenko and Ukhov (2009), and Fernández-Amador et al. (2013). We order our variables as follows: macroeconomic variables, IP, IR and MP first, followed by STD, RET and (il)liquidity. We put liquidity and illiquidity at the end of the VAR ordering in our estimates to gain stronger statistical power (see Goyenko and Ukhov, 2009).

The accumulated responses of market (il)liquidity to a unit standard deviation innovation in monetary policy shocks are summarized in Table 4, traced forward over a period of 12 months. Here, responses are measured using standard Cholesky decomposition of the VAR residuals. Bootstrap 95 percent confidence bands are used to gauge the statistical significance of the responses. The second to fifth columns (i.e. Group A) of Table 4 illustrate the aggregate response of the four (il)liquidity measures to a one-time shock in base money growth. Most of the signs are in line with our hypothesis and significant. Here, following the base money growth shocks the trading volume and turnover rate get a positive response and illiquidity ratios get a negative response. That indicates that market liquidity increases (decreases) with higher (lower) broad money supply in the economy for seven markets but not Pakistan. For Pakistan our results show that market liquidity (illiquidity) decreases (increases) with easing (tightening) of monetary policy. Similar to the money supply growth rate, the impulse response signs for the bank rate (Group B) and short-term interest rate (Group C) are found as expected, however, their magnitudes are different. In general, higher (lower) bank rate and short-term interest rate represent conservative (expansionary) monetary policy, thus generating negative (positive) influences on liquidity (illiquidity). The signs, such as bank rate impact on turnover rate in Bangladesh and Indonesia; on turnover rate and trading volume in Pakistan; and on Amihud's (2002) illiquidity ratio in Bangladesh are not found as expected. India and Indonesia also display some deviation from the expected impulse response signs of short-term interest rate shocks.

Table 4: Summary of impulse response function signs to monetary policy shocks

	Group A: Base money growth				Group B: Bank rate				Group C: Short-term interest rate			
	(il)liquidity measures											
	TR	TV	ILLIQ	TPI	TR	TV	ILLIQ	TPI	TR	TV	ILLIQ	TPI
Bangladesh	+	+	-	-	+	-	-	+	-	-	+	+
India	+	ns	-	-	-	-	+	+	+	-	-	-
Indonesia	+	+	-	-	+	-	ns	ns	+	-	ns	ns
Malaysia	+	+	-	-	-	-	+	+	-	-	+	+
Pakistan	-	-	+	+	+	+	+	+	-	-	ns	ns
South Korea	+	+	-	-	-	-	ns	ns	-	-	+	+
Taiwan	+	+	-	-	-	-	+	+	-	-	+	+
Thailand	+	+	+	+	-	+	ns	+	-	+	ns	+

Note: + and – are positive and negative responses of four (il)liquidity measures to a unit standard deviation innovation in the monetary policy variables. ‘ns’ indicates no significant positive or negative response.

From these results it is clear that, first, the monetary policy variables of Malaysia and Taiwan have the least impact on both liquidity and illiquidity compared to the other Asian equity markets considered in this study. Second, between Amihud’s (2002) illiquidity ratio and the turnover price impact of Florackis et al. (2011), the response to money growth shocks is stronger for Amihud’s (2002) illiquidity ratio. Third and finally, in line with the Granger causality results, the (il)liquidity of these eight markets is more sensitive to the short-term interest rate and the broad money supply growth rate than to bank rate. Therefore, based on the overall impulse responses, we can conclude that stock market liquidity (illiquidity) tends to rise (decline) as the base money growth increases. Conversely, market illiquidity (liquidity) tends to rise (decline) as the bank rate, cash reserve ratio and short-term interest rate increase. That means expansionary (contractionary) monetary policy of central bank increases (decreases) the liquidity (illiquidity) in Asian equity markets. The results are consistent with previous studies, such as Chordia et al. (2005), Goyenko and Ukhov (2009) and Fernández-Amador et al. (2013).

As an alternative measure recommended in many other papers including Chordia et al. (2005), we consider the variance decomposition (results available on request) of the liquidity measures to disentangle the information contributed by the monetary policy measures. The results indicate that the base money growth and bank rate respectively can explain more than a 10 percent variance of trading volume and Amihud’s (2002) illiquidity ratio in Bangladesh. Their effects also stabilize after 3 and 6 months respectively. The short-term interest rate has a weak power to explain the variance of any of the four (il)liquidity measures. For other markets, such as broad money supply and bank rate in India and Taiwan it can explain up to a 12 percent variance of TR

and TV. The short-term interest rate can explain about 13 percent of Amihud's (2002) illiquidity and TPI ratio. Similarly, the short-term interest rate can explain the variance of TV and Amihud's (2002) illiquidity ratio up to 10 and 7 percent respectively in Indonesia and South Korea; bank rate can explain up to 8 and 12 percent of Amihud's (2002) illiquidity ratio in Malaysia and Pakistan respectively; and bank rates and short-term interest rates can explain up to 11 percent in Thailand. For the other macroeconomic variables, we find that both inflation rate and industrial production have a significant effect on market liquidity. Own-market volatility and returns are other important variables as suggested in Chordia et al. (2005) that can significantly explain the variance of the market (il)liquidity and in our sample the maximum influence is up to 20 percent.

3.2.2 The impact of fiscal policy shocks

For fiscal policy and liquidity measures, we run 12 different VAR models for each equity market and the pairwise Granger causality between endogenous variables associated with those VAR models. Here, the null hypothesis we test is that the lagged endogenous variables of interest (i.e. fiscal policy or liquidity measures) do not Granger cause the dependent variable of interest (again, either market liquidity or fiscal policy variables). The results shows that fiscal policy variables have causal and reverse-causal associations with both liquidity and illiquidity measures (detailed results are available on request). In particular, government expenditure has significant impact in Bangladesh, India and South Korea; government borrowing in Bangladesh, India and Taiwan; and public borrowing in India and Pakistan. For example, government expenditure has a significant causality with trading volume in Bangladesh, Indonesia and South Korea. Similarly, Amihud's (2002) illiquidity measure is Granger caused by government borrowing and public borrowing in Bangladesh, India, Pakistan, Taiwan and Thailand; the coefficients are statistically significant at 5 and 10 percent respectively. However, we see there is only reverse causality between the turnover rate and private borrowing in Bangladesh and Malaysia. Altogether, these results support the long-run cointegration found between liquidity measures and fiscal policy variables in section 3.1.

Our results further show a strong two-way causation between fiscal policy and other endogenous variables of the VAR. For example, government expenditure has a one-way causation with market returns in Bangladesh and this association is statistically significant at a 5 percent level. Government borrowing and public borrowing have significant causality with returns in India and Indonesia. The bidirectional causality identified in India, Indonesia, Pakistan, South Korea, Taiwan and Thailand is, however, mostly with government borrowing. Besides, market volatility

is significantly Granger caused by government expenditure and government borrowing in Bangladesh, Indonesia and Pakistan. A reverse impact exists with government borrowing in Indonesia, South Korea and Thailand; and with private borrowing in Taiwan. All these bidirectional causations of stock returns and volatility imply that sometimes market characteristics can influence the government's fiscal decision. There is empirical evidence available on this relationship. For example, Tagkalakis (2011) reports that changes in equity price have an impact on government expenditure and revenue.

Similar to the monetary policy influences and the recommendation of Gagnon and Gimet (2013), we report the impulse response functions and variance decomposition to better understand the dynamics of fiscal policy within the VAR system. We order our variables as follows: macroeconomic variables - IP, IR and FP first, followed by STD, RET and (il)liquidity ratios. The signs of accumulated responses of market liquidity to a unit standard deviation innovation of fiscal policy shocks are presented in Table 5 (from Groups A to C), traced forward over a period of 12 months. In each group the four variables represent the two liquidity and two illiquidity responses to fiscal policy variables. The responses are estimated using a standard Cholesky decomposition of the VAR residuals and use the bootstrap 95 percent confidence bands to gauge the statistical significance of the responses.

Table 5: Summary of impulse response function signs to fiscal policy shocks

	Group A: Government Consumption				Group B: Government Borrowing				Group C: Private Borrowing			
	(il)liquidity measures											
	TR	TV	ILLIQ	TPI	TR	TV	ILLIQ	TPI	TR	TV	ILLIQ	TPI
Bangladesh	+	+	-	+	-	-	+	+	+	+	+	+
India	+	+	-	-	-	-	+	+	+	+	-	-
Indonesia	+	+	-	-	-	-	+	+	+	+	-	-
Malaysia	+	+	-	-	+	+	-	-	+	+	-	-
Pakistan	+	+	-	+	+	+	-	+	+	+	-	+
South Korea	ns	+	+	-	-	-	+	+	+	+	-	-
Taiwan	+	+	-	-	-	-	+	+	+	+	-	-
Thailand	+	-	-	+	+	-	+	+	+	+	-	-

Note: + and - are positive and negative responses of four (il)liquidity measures to a unit standard deviation innovation in the monetary policy variables. 'ns' indicates no significant positive or negative response.

The accumulated responses (from Group A to C) show that market liquidity increases following government expenditure and private borrowing shocks, and decreases with government borrowing shocks. Surprisingly, some of the markets do not show the usual reaction to fiscal policy shocks, such as in Pakistan. The illiquidity ratio (i.e. turnover price impact) of this market

gives a positive response to government consumption and private borrowing innovations. Furthermore, government borrowing and market liquidity ratios are positively related in both Pakistan and Malaysia. Similarly, the response of turnover price impact in Bangladesh, Pakistan and Thailand; the response of Amihud's (2002) illiquidity in South Korea; and the response of turnover rate in Thailand are positive. Other than these, most of our impulse response signs (i.e. eighty signs out of ninety six) are found as expected and explained in theory. In particular, trading volume and turnover rate increases over the 12 month period due to any changes in government expenditure or private borrowing. The influence is strongly more positive for private borrowing than government expenditure on illiquidity ratios in India, Malaysia, Pakistan and Thailand. On the contrary, trading volume more significantly increases due to any shocks from government expenditure than private borrowing in Bangladesh, Indonesia, South Korea and Taiwan. Therefore, all together, we can state that expansionary (contractionary) fiscal policy increases (decreases) the market liquidity of the Asian equity markets. In this vein, it is documented in earlier literature that fiscal shocks or a large budgetary stimulus can facilitate access to credit for the private sector (see Blanchard, 2009; Gagnon and Gimet, 2013) and this credit could channel to the stock market to boost liquidity.

On the other hand, as expected, government borrowing significantly influences Amihud's (2002) illiquidity ratio and turnover price impact (see Group B in Table 5). In addition, signs of impulse response functions show that government borrowing reduces the liquidity related ratios except in Malaysia and Pakistan (however, turnover price impact ratio increases with government borrowing). Combined, these results support a 'crowding out' effect in this economy created by the government and thus business firms suffer from a lack of credit opportunity. Interestingly, we also find a strong effect of government expenditure and private borrowing on illiquidity measures of India, South Korea and Thailand.

In line with monetary policy and the arguments of earlier papers, we estimate the variance decomposition of liquidity measures associated with fiscal policy variables (not reported but available on request). We can make several conclusions from these results. First, government expenditure and private borrowing have greater power than government borrowing to explain the variation in liquidity and illiquidity. For example, government expenditure (private borrowing) can explain up to 12 (10) percent of the variance in turnover rate and up to 15 (12) percent in trading volume. Government borrowing, on the other hand, can contribute around 8 percent of information to Amihud's (2002) illiquidity ratio and around 6 percent of information to the turnover price impact of Florackis et al. (2011). Second, at a short horizon private

borrowing significantly increases market liquidity by up to 10 percent, however, it stabilizes after six months. Third, for most markets, the percentages explained by fiscal policy measures stabilize after 3 to 6 month periods. Fourth, the effect of fiscal policy is comparable to or even larger than other macroeconomic variables, such as industrial production and inflation rate. However, own-market returns and standard deviation are often important factors for explaining the liquidity variances. Fifth, for some markets, such as Bangladesh, government borrowing is a significant determinant of market (il)liquidity. Finally, the amount of information contributed by fiscal policy in liquidity variance is higher than the monetary policy variables for some countries, such as Bangladesh and Pakistan.

Interestingly, it is documented from the results of impulse responses and variance decomposition that the liquidity measures of Asian stock markets are not only influenced by monetary policy but fiscal policy variables also have significant effect. Nevertheless, there is little empirical evidence available in this area of fiscal policy except Gagnon and Gimet (2013). In their paper, Gagnon and Gimet (2013) report that government spending has a positive impact on credit and consumption. Earlier studies, e.g. Spilimbergo et al. (2009), Blanchard et al. (2010), and Eggertsson and Krugman (2012) support this idea and suggest that tax breaks, fiscal stimulus and expansionary fiscal policy can increase firms' and investors' access to credit. In particular, Spilimbergo et al. (2009) assert that tax breaks can strengthen consumers' and companies' financial health, which would in turn increase their access to credit and thus enhance market liquidity.

Beside this limited empirical documentation on the influence of fiscal measures on liquidity, there is much evidence available on the connection between fiscal policy and other characteristics of financial markets from both developed and emerging economies. For example, Darrat (1988) finds that fiscal deficit exerts a highly significant negative impact on the current stock market of Canada. Using a panel of OECD countries, Ardagna (2009) reports that fiscal adjustments based on expenditure reductions are related to an increase in stock market prices. Montes and Tiberto (2012) provide evidence from Brazil suggesting that the efforts of both fiscal and monetary authorities have been essential for macroeconomic stability, and thus to stimulate the stock market. Similarly, from a multi-country data set Chatziantoniou et al. (2013) report that both fiscal and monetary policies influence the stock market, via direct or indirect channels. Recently, Belo et al. (2013) claim government spending has an impact on expected firm cash flows. In addition, uncertainty about the impact of government policies can affect the rate at which future cash flows are discounted. Pástor and Veronesi (2012) comment that

governments shape the environment in which the private sector operates and stock prices should fall on the announcement of policy changes, on average.

Sometimes the impact of monetary policy depends on fiscal policy, e.g. Jansen et al. (2008) maintain that the effect of monetary policy on the stock market varies, depending on the fiscal policy stance. Moreover, when the equity market is influenced by both policies, as happens in most of our sample Asian markets, the direction of final interaction between these two variables is important. This is because when the central bank formulates disinflationary policies while the government is engaged in expansionary strategies then the ultimate outcome will deviate significantly from the desired objective (see Dixit and Lambertini, 2003). For investors our findings have significant implications. They have to understand the interaction between macroeconomic variables and the characteristics of a stock market while making their investment decisions. More specifically, they should consider the cyclic association between fiscal and monetary policy rather than the isolated impact on the stock market.

3.2.3 The impact of the 2007-8 financial crisis

It is documented in earlier literature that during a financial crisis market conditions can be severe and liquidity can decline or even disappear. As Chordia et al. (2005) suggest, such liquidity shocks are a potential channel through which asset prices are influenced by liquidity. In this vein, Næs et al. (2011) mention a possible causal link between a decline in the liquidity of financial assets and economic crises. Many other empirical papers, such as Amihud et al. (1990), Liu (2006), Brunnermeier and Pedersen (2009), Rösch and Kaserer (2013) and Gagnon and Gimet (2013) have highlighted the relationship between equity market liquidity and various financial crises. We, following these earlier studies, investigate the influence of the financial crisis (i.e. 2007-08) on the liquidity measures of these eight emerging stock markets of Asia. Since the crisis of 2007-08 expanded and became global, emerging markets also began to experience liquidity strains (Yehoue, 2009). Similarly, Blanchard (2009) mentions that by 2008 the credit freeze crisis had spread internationally, causing a dramatic global decrease in stock prices and a fall in consumers' and firms' confidence. Previously, on emerging markets liquidity, Lesmond (2005) and Yeyati et al. (2008) have examined the impact of financial crises. However, Lesmond (2005) shows the impact of the Asian and Russian crises and Yeyati et al. (2008) consider the crisis episodes over the period from April 1994 to June 2004.

Firstly, we tested for Granger causality between market liquidity measures and the financial crisis (results available on request). The results indicate that the 2007-08 crisis has significant influence

on market illiquidity. For example, the crisis has a strong causal effect on Amihud's (2002) illiquidity ratio in Indonesia, Malaysia, Pakistan and Taiwan. Similarly, there is a relationship between the turnover price impact ratio and the crisis in Malaysia, Pakistan and Taiwan. Furthermore, Malaysia has a bidirectional causal effect from illiquidity measures. Due to this result, we cannot deny the argument of Amihud and Mendelson (1986) and Næs et al. (2011) that illiquidity does not lead to a crisis, or a decline in the liquidity of financial assets does not Granger cause an economic crisis. On the other hand, the crisis of 2007-08 Granger causes the liquidity ratios in Bangladesh, India, Indonesia and Malaysia. For example, there is one-way causality with turnover rate and trading volume in Bangladesh and with turnover rate in India. In particular, we can accept the alternative hypothesis that a crisis Granger causes turnover rate at the 1 and 7 percent levels respectively for Bangladesh and India; and at the 10 percent level for trading volume in Bangladesh. There is reverse causality between a crisis and the turnover rate in Indonesia and Malaysia. That means that during the financial crisis the trading activities of Asian equity markets were significantly affected. A similar influence of a financial crisis on trading activity was identified by Yeyati et al. (2008) in their emerging markets sample.

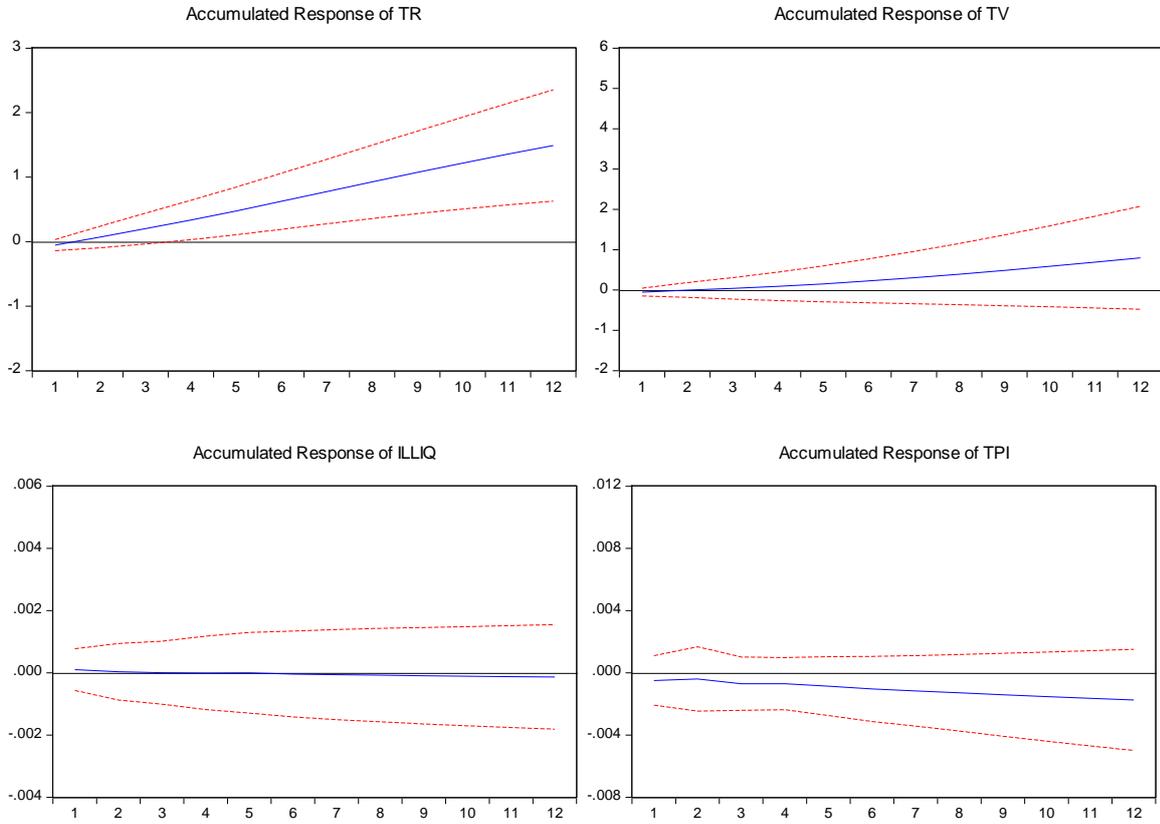
The impulse responses of our VAR system are reported in Figure 3 (from Group A to Group H). Figure 3 shows the accumulated responses of market liquidity to a unit standard deviation innovation in a crisis, traced forward over a period of 12 months. Similar to previous sections (i.e. 3.2.1 & 3.2.2), the responses are estimated using a standard Cholesky decomposition of the VAR residuals and the bootstrap 95 percent confidence bands gauge the statistical significance of the responses. Surprisingly, our results indicate that Asian markets received differential impacts from the crisis of 2007-08. In most of the markets (i.e. India, Indonesia, Malaysia, Pakistan, and Taiwan) illiquidity related measures were affected positively, where Amihud's (2002) illiquidity ratio (in Indonesia, Malaysia, Pakistan and Taiwan) received a greater impact than the turnover price impact (significant in India and Taiwan). This finding supports the conventional argument mentioned in Yeyati et al. (2008) that during a crisis, prices change more with each dollar transacted, pushing Amihud's illiquidity measure up, and bid-ask spreads widen. The drying-up of market liquidity during the financial crisis is a well-documented phenomenon, responsible for the financial contagion experienced during the crisis (see Rösch and Kaserer, 2013).

Instead, we see both of the liquidity related measures (i.e. turnover rate and trading volume) increase following the crisis shock in Bangladesh, and only the trading volume in India. For Bangladesh an innovation to the crisis is strongly significant on the turnover rate with the response peaking from period one. However, the trading volume starts to increase from the

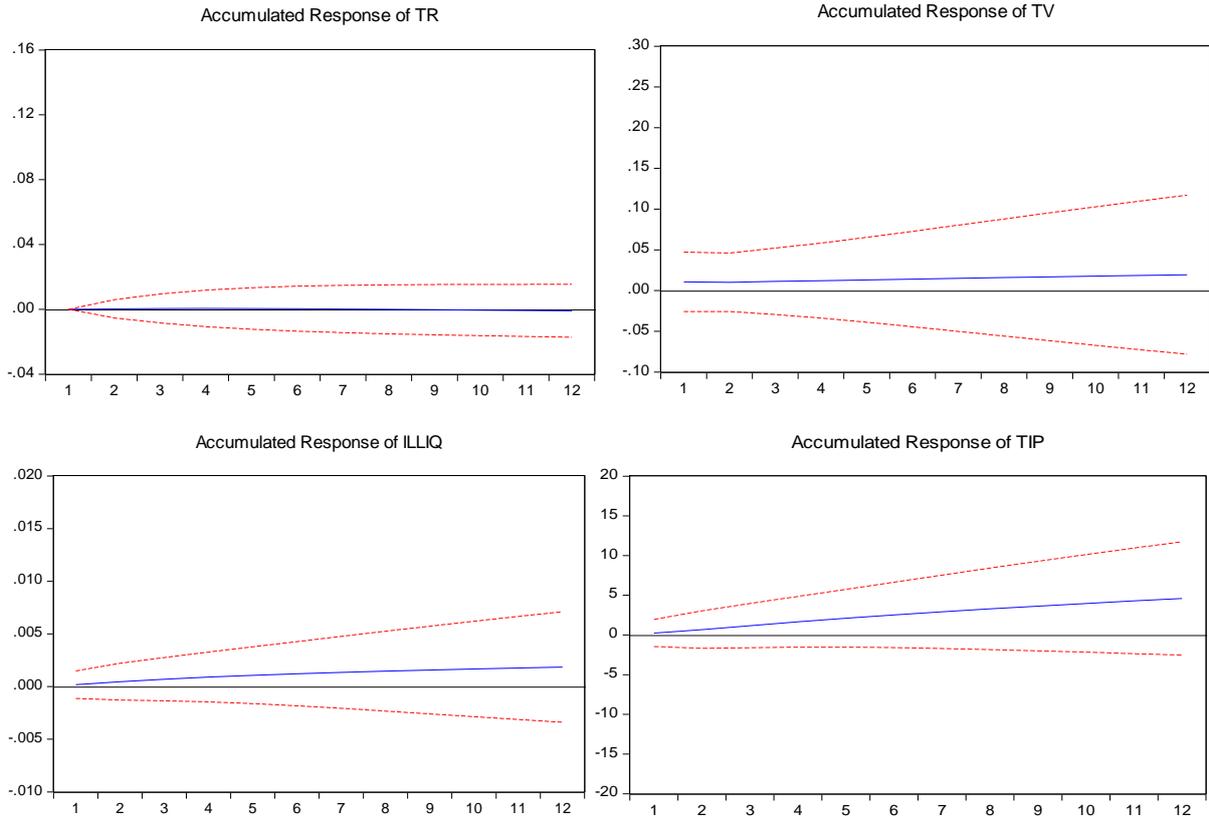
fourth month following an innovation to the crisis. Except these two countries, trading volume receives significant negative shocks in South Korea, Taiwan and Thailand; and turnover rate receives negative shocks in Indonesia, Malaysia, Pakistan, South Korea and Thailand. Our results therefore highlight the fact that the nature and depth of the financial crisis of 2007-08 was different from other financial crises, or the crises that happened in certain regions, as suggested in many earlier studies, such as Bartram and Bodnar (2009). For example, Lesmond (2005) analyses 23 emerging markets and finds that bid-ask spreads and several other liquidity measures sharply increased during the Asian and Russian financial crises. Yeyati et al. (2008) also find an initial increase in trading activity during the crisis period by examining seven different countries and 52 different stocks over a crisis period from April 1994–June 2004. On the contrary, due to an absence of liquidity during the crisis of 2007-08, the world faced dramatic falls in stock markets and all over the world central banks and governments adopted massive stimulus packages to increase liquidity (see Söderberg, 2008; Blanchard et al., 2010; Woodford, 2011). Global equity market capitalization fell from \$51 trillion to \$22 trillion over the 17 month period from October 2007 to February 2009 (see Bartram and Bodnar, 2009). In particular, the onset of the global financial crisis dealt a heavy blow to highly export dependent Asian economies, such that East Asia's capital account suffered a \$172 billion fall in just a year from 2007 to 2008; Asian markets experienced a \$35.7 billion decline from their portfolio investment account in 2007; and in 2008 the net outflow from the emerging Asian markets took up 61.3 percent of total net outflow from the emerging markets (see Park, Pempel and Xiao, 2012 for more discussion).

Figure 3: Accumulated response of (il) liquidity variables to Cholesky One S.D. crisis shocks

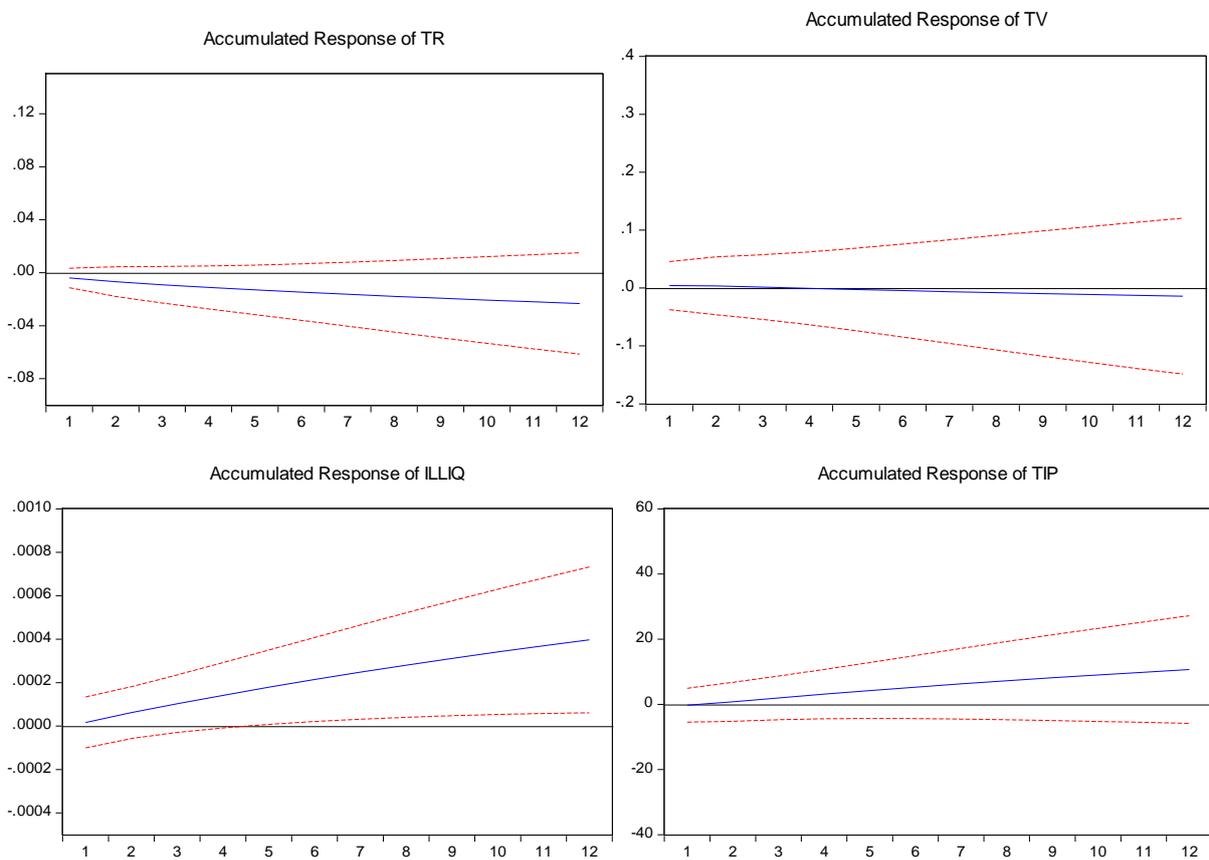
Group A: Bangladesh



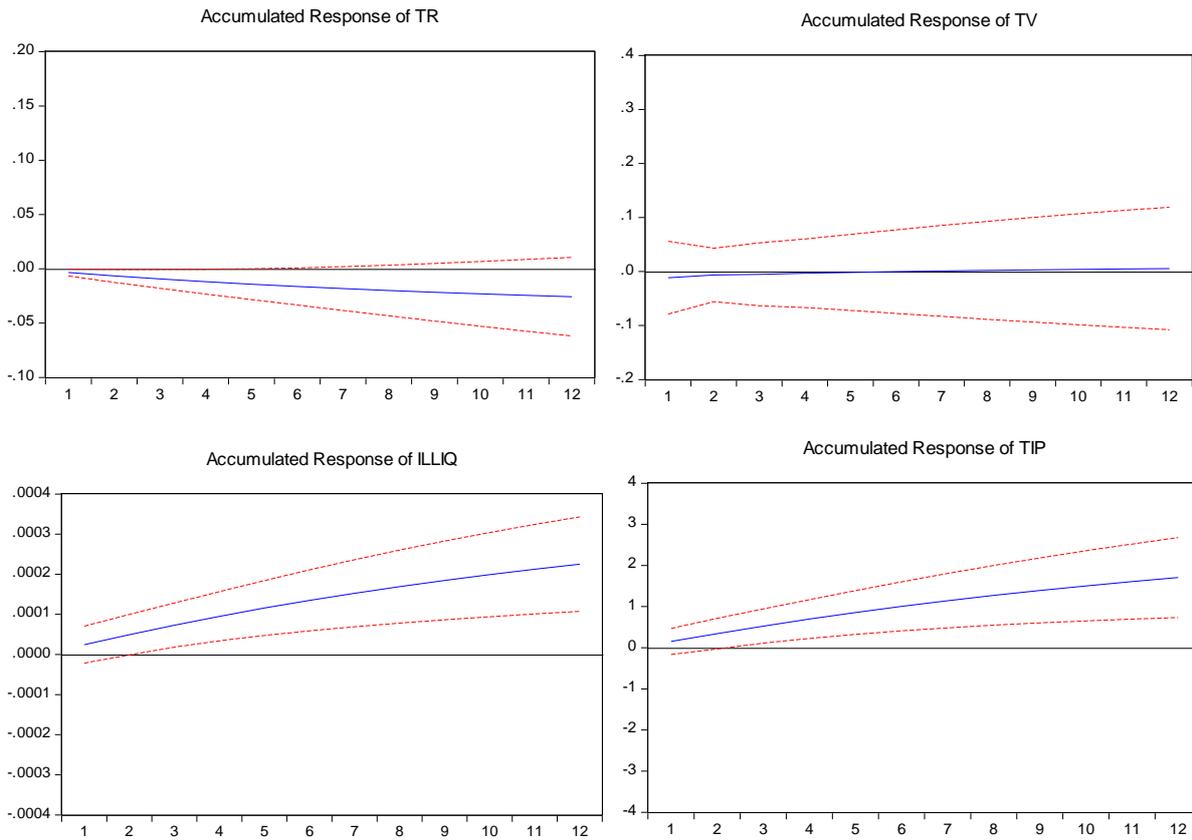
Group B: India



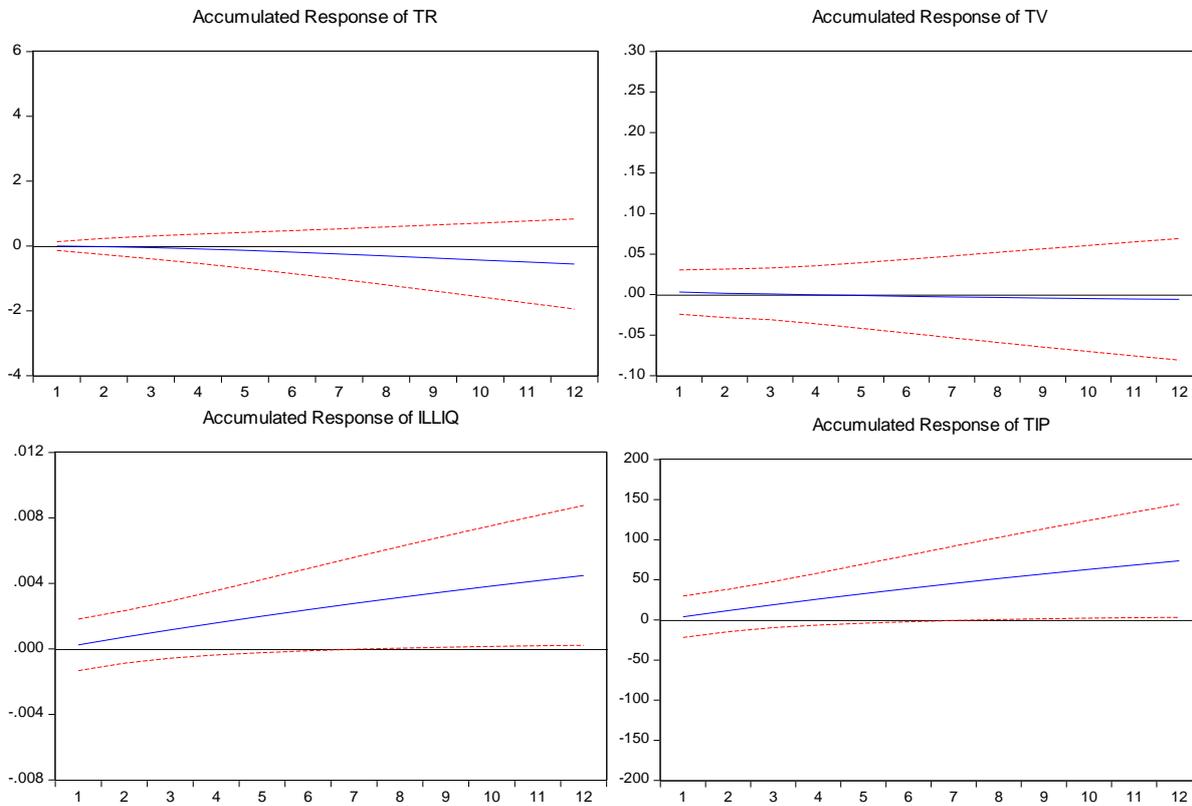
Group C: Indonesia



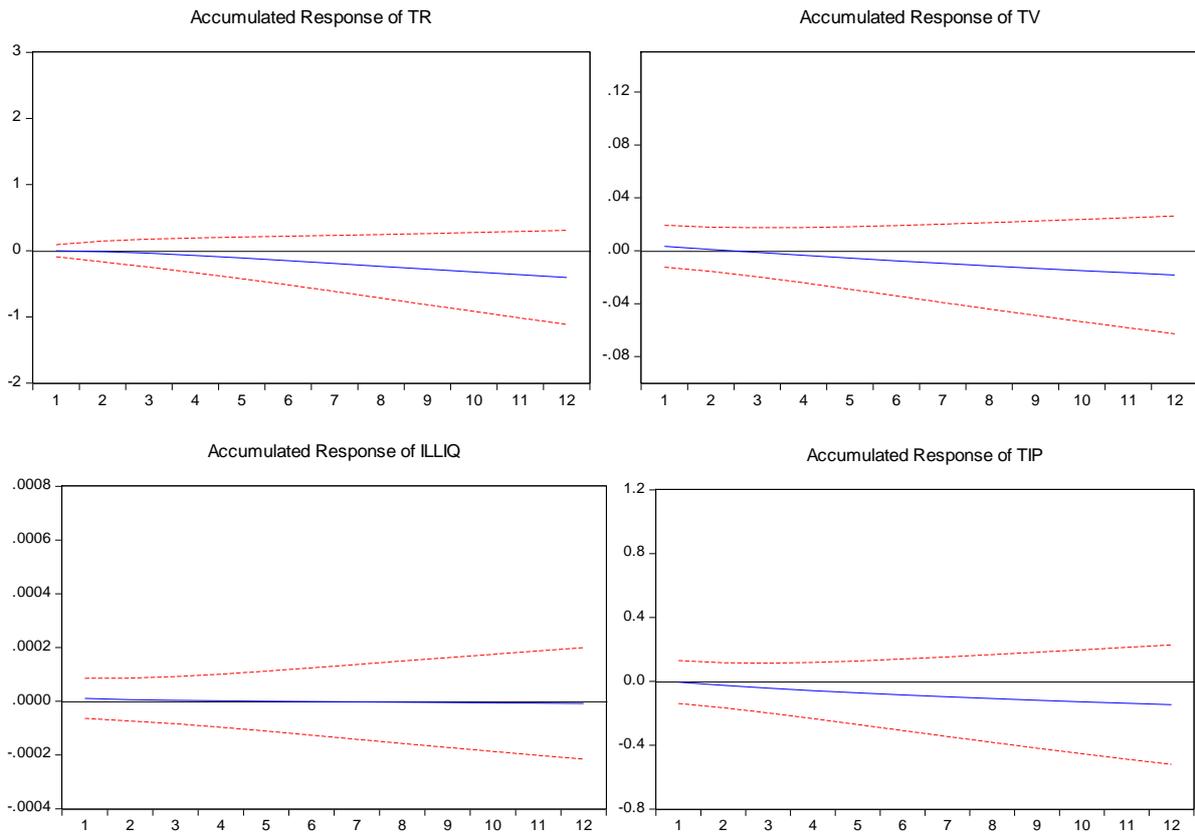
Group D: Malaysia



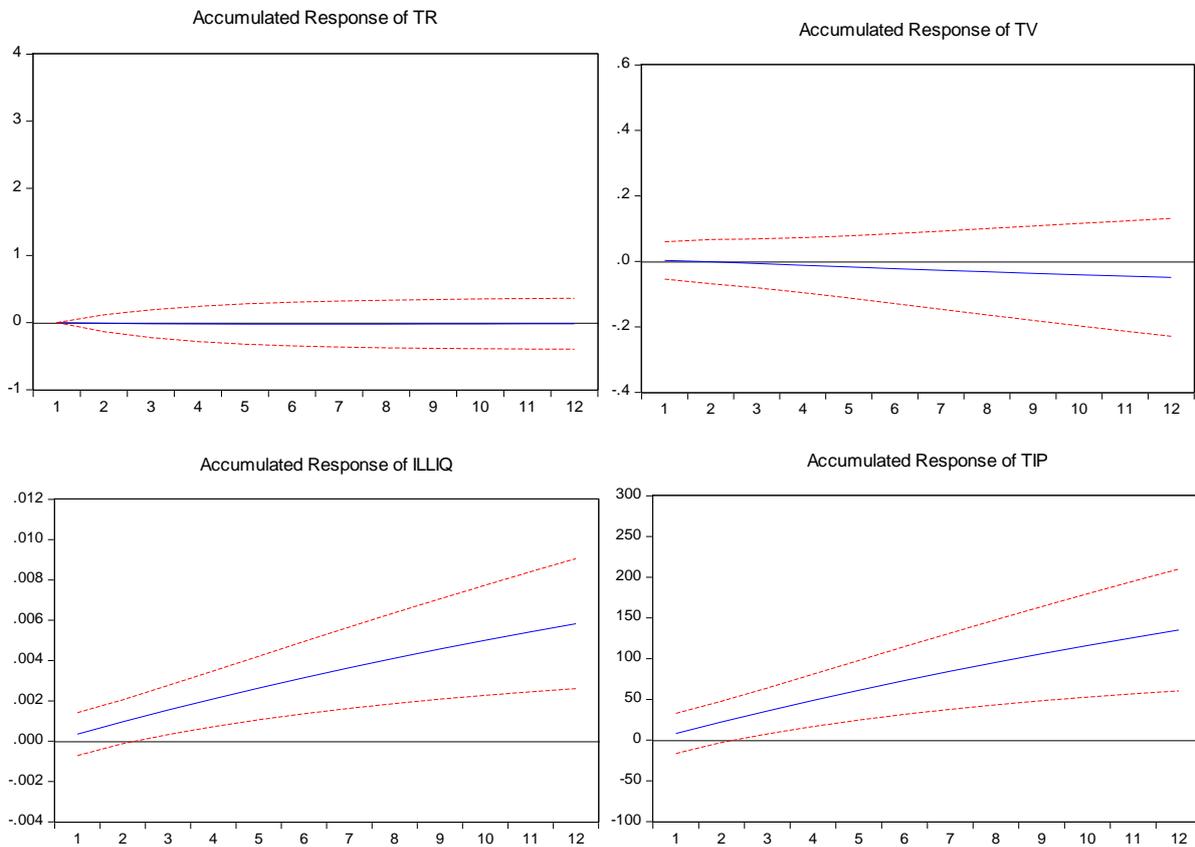
Group E: Pakistan



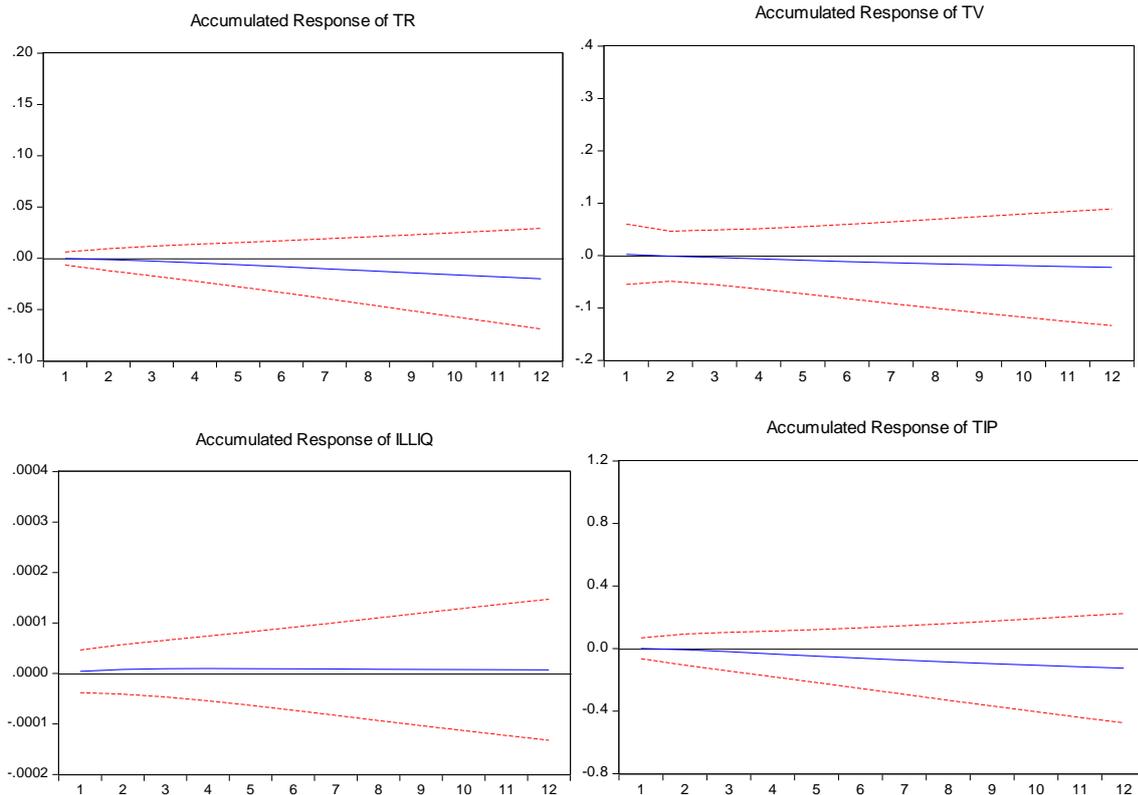
Group F: South Korea



Group G: Taiwan



Group H: Thailand



To get a clear and comparable picture of the influence of the crisis of 2007-08, we ran the variance decomposition for the crisis periods and checked the impulse response function for non-crisis periods. The result of the variance decomposition shows that the financial crisis can explain around 27, 25, 15 and 5 percent of the error variance of Amihud's illiquidity ratio, turnover rate, turnover price impact and trading volume respectively. On the other hand, results of a VAR impulse response function (signs of non-crisis shocks are reported in Table 6) indicate that most of the Asian markets in our sample generally follow the conventional phenomenon related to (il)liquidity ratios during non-crisis periods. That is, Amihud's illiquidity ratio and turnover price impact receive significant negative shocks and liquidity ratios receive positive shocks. Hence, these results support our argument that the crisis of 2007-08 had different depth and breadth compared to other crises and now equity markets are more integrated than in the 1980s or 1990s. However, surprisingly, Bangladesh has negative and positive shocks respectively on liquidity and illiquidity during this non-crisis period. There could be two explanations for this: (i) this market is not influenced by international factors, rather local macroeconomic factors are more important (ii) due to this lower integration with the world market, Bangladesh received \$1.09 billion of foreign direct investment (FDI) during 2007-08. Moreover, due to local political stability the equity market has increased by 66% in 2007-08. As reported in *The Economist*

(2011), since 2007, the Bangladesh stock market outperformed almost all the world's markets, gaining as much as 410% in value over the period to 2010.

Table 6: Summary of impulse response function signs to non-crisis shocks

	Bangladesh	India	Indonesia	Malaysia	Pakistan	South Korea	Taiwan	Thailand
TR	-	+	+	ns	+	+	-	+
TV	-	ns	ns	ns	ns	+	+	ns
ILLIQ	+	-	-	-	-	ns	-	ns
TPI	+	-	-	-	-	+	-	+

Note: + and – are positive and negative responses of four (il)liquidity measures to a unit standard deviation innovation in the monetary policy variables. ‘ns’ indicates no significant positive or negative response. To understand the relationship from a deeper perspective, we further analyse the macroeconomic influence on market liquidity by selecting a calm episode, 2001:Q3 – 2004:Q4. Results of our calm episode are found to be mostly similar to the aggregate non-crisis reported in this Table 6. Interestingly many of the ‘ns’ become significant but signs are found as expected, such as positive (+) sign of TR for Malaysia and positive (+) of TV for India, Indonesia and Pakistan. The only sign that has changed in a calm episode compared to our non-crisis period is the sign of TR for Taiwan. However, that sign has changed to expected (i.e. +) from unexpected (i.e. -).

3.3 Firm and industry level evidence

In line with our second objective this section provides evidence related to the influence of monetary and fiscal policy variables on the liquidity of individual firms traded in our eight Asian stock markets. Using equations (vi) and (vii), in the first step, we run panel regressions on individual firms and then in the second step, panel regressions on major industries to see the impact of macroeconomic management policy. The outcomes are reported in Tables 7 to 10. The interaction term ($Int_{i,t-1}$) in our models for individual firms indicates whether the influence of monetary (MP) or fiscal policy (FP) depends on firm size as measured by logged market capitalization. However, there is no such term in our second set of models, i.e. panel estimation for industries (Table 10), since firms are not sorted by size, but by sector. We test for stationarity applying the panel unit root test developed by Levin et al. (2002) and Pesaran (2007) and use first difference where we failed to reject the null of a unit root, which further minimizes the multicollinearity problem. In order to account for heteroskedasticity, all p-values are based on robust standard errors. The lag order for each variable is selected based on whether any autocorrelation exists in the residuals.

3.3.1 Macroeconomic management and individual firms

We estimate panel regressions for each of the four (il)liquidity measures and the four monetary and three fiscal policy variables for each market. Table 7 presents the influence of monetary policy and other macroeconomic variables on (il)liquidity measures. The impact of fiscal policy

as well as other control variables on (il)liquidity is reported in Table 8. Due to the use of an interaction term, it is convenient to evaluate the effects of the monetary and fiscal policy at different percentiles of the sample distributions for the interaction variable (market capitalization) (see Fernández-Amador et al., 2013). The impact of monetary and fiscal policy is evaluated at the minimum 20%, 40%, 60%, 80% and maximum. Here, we report the sign, magnitude and significance of the interaction term for each of these size-based portfolios in Table 9. Each table represents the marginal effects at the average values of the distribution and the p-value is estimated based on average t-statistics.

The second to fifth columns of Table 7 show the estimation results when measuring monetary policy by the rolling twelve-month growth rate of base money. Results are significantly positive for turnover rate and trading volume. Thus, on average an increase in growth rate of base money leads to a rise in trading activity, while illiquidity declines in Asian emerging markets. This further implies that an expansionary (contractionary) monetary policy increases (decreases) individual stocks' liquidity. However, our particular interest is to see the results of interaction terms, i.e. whether the impact of money supply growth depends on size of firms. The estimations show (Table 9) that the liquidity measures (i.e. TR and TV) of smaller firms (i.e. bottom 20%) and illiquidity ratios (i.e. ILLIQ and TPI) of larger firms (i.e. top 20%) are significantly influenced by the money supply growth rate. However, the average coefficient of Amihud's (2002) illiquidity ratio for the smallest 20 percent of firms (-0.594) is larger in absolute size than the same average coefficient of the largest 20 percent of firms (-0.0297) in Asian equity markets.

The rest of the results in Table 7 are related to panel estimations where central bank policy is approximated by the bank rate and short-term interest rate. We see in each model the coefficients of monetary policy are negative for liquidity measures and positive for illiquidity measures, except the average sign of the turnover price impact coefficient to short-term interest rate shocks. In particular, the impact of bank rate is positive and statistically significant on Amihud's (2002) illiquidity ratio and turnover price impact in Bangladesh. However, a positive influence of bank rate on trading volume is identified in India, Malaysia, Taiwan and Thailand. Similarly, the short-term interest rate has a significant negative influence on turnover rate in India, Bangladesh, Indonesia, Taiwan and Thailand. These results are well in line with our hypothesis that any such restrictive monetary policy decreases the liquidity and increases illiquidity for the overall equity market. The interaction term (Group A of Table 9) generally confirms the empirical pattern observed in base money growth with slight differences. On average the top and bottom 20 percent portfolios are more sensitive to monetary policy and the

illiquidity and liquidity ratios respectively are highly influenced by those policy variables. We find that an increase in bank rate and short-term interest rate increases (decreases) the liquidity of larger firms (smaller firms). That means that during restrictive monetary policy investors prefer larger firms to smaller. However, individually, for Indonesia, Pakistan and Thailand we see a greater impact of money supply and bank rate changes at firm level, whereas in India and South Korea firms are largely affected by any changes in the short-term interest rate. Finally, it is interesting that regardless of the size of firms, the trading volume ratio is influenced by all four central bank policies and their coefficients are also statistically significant.

Remarkably, Table 9 (Group A) provides results from emerging Asian markets in line with many developed equity markets. For example, for the NYSE, Amihud (2002) reports that small stocks are more responsive to illiquidity shocks, while large stocks become more attractive when aggregate liquidity declines. That means that expansionary policy reduces the illiquidity for smaller firms significantly more than larger firms and thus during a restrictive period larger firms become preferable. On average a similar trend is reported in our results. We find that the coefficient of monetary base for the bottom 20 percent of firms are +0.91 (TR), +2.979 (TV) and -0.594 (Amihud's illiquidity ratio). That means that the liquidity (illiquidity) of smaller firms increases (decreases) with a greater supply of money in the economy. On the other hand, Næs et al. (2011) and Fernández-Amador et al. (2013) also assert that smaller firms are more responsive to liquidity shocks, which implies contractionary policy has a greater impact on smaller firms. Besides, we would expect the liquidity variation of small firms to be higher than the liquidity variation of large firms (Næs et al., 2011). In particular, from the German, French and Italian stock markets, Fernández-Amador et al. (2013) report that the liquidity-providing effect of a loose monetary policy (i.e. higher money supply) is stronger for larger firms, but the impact of restrictive monetary policy (i.e. higher EONIA) tends to decrease with increasing firm size both for liquidity and illiquidity measures. In this line of thought, Næs et al. (2011) comment that variation in market liquidity is caused by portfolio shifts, from illiquid and more risky assets to safer and more liquid assets, due to changing expectations about economic fundamentals or binding funding constraints.

Table 7: Panel estimation for monetary policy impact on individual firms

	Monetary Base				Bank Rate				Short-term interest Rate			
	TR	TV	ILLIQ	TPI	TR	TV	ILLIQ	TPI	TR	TV	ILLIQ	TPI
Dependent variable _{t-1}	0.6426***	0.7749***	0.3221***	0.3243***	0.6623***	0.7723***	0.3172***	0.3209***	0.6007	0.7632***	0.2218	0.2762***
Monetary policy _{t-1}	0.0048***	0.0555***	-1.1E-05	-0.1418	-0.0831	-0.1639**	0.00012	1.3697	-0.0551***	-0.0300	7.2E-05	-0.4312***
Interaction term _{t-1}	0.0111*	0.3446***	-9.7E-08	0.0175	1.98E-05	0.0225***	-1.1E-05	-0.0460***	0.0005	0.0149	-2.5E-06	0.0019
Return _{t-1}	3.9430	1.8617	-0.0003	8.2318	3.6623	1.8039	-0.00023	15.1862	3.2929	1.7687	0.0001	8.0419
Standard deviation _{t-1}	-3.6139*	-4.4822*	0.0004	-8.9091	-3.4418*	-4.3091*	0.0007	-18.9653*	-3.5044	-4.7034*	-0.0003	-9.7362
Ln(Market value) _{i,t-1}	-0.0362	-0.0439***	1.85E-05	-0.0490	-0.0353	0.0164**	-3.1E-05	0.3046	-0.0341	0.0146***	0.0004**	-0.0109
Industrial production _{t-1}	0.0020***	0.0047*	2.67E-08	-0.1896	0.0019***	0.0088***	6.25E-07	-0.5476*	0.0018***	0.0034	2.69E-06	-0.1894
Inflation _{t-1}	-0.0190	0.0221	8.69E-05	0.2089	-0.0313	-0.00807	8.21E-05	-0.0128	0.0299	0.0655	4.31E-05	0.3960
Stock index _{t-1}	0.1138	0.1223	3.63E-05	-0.2385	0.0893	0.0739	5.41E-05	-0.3654	-0.0139	0.2868	-0.0011	-0.8441***
R	0.6126	0.8466	0.2911	0.2646	0.6276	0.8496	0.2970	0.2805	0.6249	0.8798	0.3167	0.2953

***, ** and * significant at the 1%, 5% and 10% levels based on average t-statistics

Overall, we see a similar shift in portfolio holdings in our sample Asian markets as suggested in Amihud (2002) and Næs et al. (2011). Investors prefer more liquid stocks during restrictive monetary policy (i.e. with higher bank and short-term interest rates) and broad money supply has a significant influence on firms' (il)liquidity. In particular, the top 40 and mid 20 percent of firms on average and firms listed in India, Indonesia, Pakistan, Taiwan and Thailand receive a higher influence of monthly changes in money supply. Moreover, since smaller firms are more dependent on bank lending than larger firms thus a restrictive monetary policy (i.e. higher short-term interest rate) may increase the cost of funds for them to a greater degree (see Fernández-Amador et al., 2013). However, the influence of central bank policy on the (il)liquidity of these emerging markets largely depends on the instruments used, and therefore investors and policy makers must be careful in measuring their impact on the overall market.

Table 8 depicts the results of panel estimations of (il)liquidity measures where fiscal policy is approximated by government expenditure, government borrowing and private borrowing. Results indicate the liquidity of individual firms is significantly and positively influenced by government expenditure and private borrowing, but negatively by government borrowing. These results are consistent with our hypothesis; expansionary (contractionary) fiscal policy by the government increases (decreases) market liquidity at the firm level. For interaction terms (Group B of Table 9), we find a similar response to that of monetary policy – both liquidity and illiquidity of large and smaller stocks are more sensitive to fiscal information. In addition, each policy variable creates a differential impact on (il)liquidity measures, however, trading volume is affected significantly by all three fiscal policies across all firm sizes, particularly in Bangladesh, India, Indonesia, Pakistan and South Korea.

From Table 9, we see that on average government expenditure increases trading activity (both turnover rate and trading volume ratio) for all size portfolios, except that smaller firms are more sensitive to fiscal policy shocks. Their liquidity rises and illiquidity (e.g. Amihud's illiquidity ratio) declines significantly. We find consistent results of government borrowing with our hypothesis as well. Liquidity (illiquidity) of both smaller and larger firms reduces (increases) due to borrowing by the government from the banking sector except in Pakistan and Malaysia. However, the magnitude of the impact is slightly higher for the smallest portfolios (i.e. the bottom 20 percent). This result supports the argument of Fisher (1988) and Fernández-Amador et al. (2013). When government borrows from the banking sector, business firms suffer and since smaller firms depend more on bank credit, they suffer more than others do.

Table 8: Panel estimation for Fiscal policy impact on individual firms

	Government Expenditure				Government Borrowing				Private Borrowing			
	TR	TV	ILLIQ	TPI	TR	TV	ILLIQ	TPI	TR	TV	ILLIQ	TPI
Dependent variable _{t-1}	0.6299***	0.7437***	0.2841***	0.3495***	0.6406***	0.7451***	0.5774***	0.3479***	0.6296***	0.7444***	0.2836***	0.3492***
Fiscal policy _{t-1}	-0.0036	0.0203**	-0.0002	-0.0026	-0.0034	0.0065**	0.0116	0.2832	0.0005	0.0098**	-0.0009	0.1216
Interaction term _{t-1}	-0.0010*	0.0129	-0.0003	0.0005	-0.0003	0.0091**	0.0009	0.0134	-0.0004	0.1971	-0.0005	0.0097***
Return _{t-1}	6.5339	1.6404	-0.1057	8.1050	6.4863	1.5560	-0.1203	10.9899	6.6914	2.1499	-0.0922	8.3217
Standard deviation _{t-1}	-3.5484*	-6.1179*	-0.1491*	-8.7595	-3.5518*	-6.1279**	0.1994	-9.7341	-3.5511**	-6.1924*	-0.1521	-9.3124
ln(Market value) _{i,t-1}	-0.0329	-0.0375***	-0.0005	-0.0446	-0.0333	-0.0518	-0.0391	-0.2025	-0.0323*	-0.0255***	-0.0004	-0.2951
Industrial production _{t-1}	0.0013***	0.0066	-0.0003	-0.1922	0.0010	0.0051	0.0031	-0.2846	0.0012***	0.0043	-0.0002	-0.1971*
Inflation _{t-1}	-0.0205	-0.0091	-0.0009	0.1982	-0.02167	-0.0208**	-0.0377	0.3653**	-0.0232	-0.0234***	-0.0010	0.2953*
Stock Index _{t-1}	0.1409	0.2061	-0.0081	-0.2339	0.1406	0.2329	-0.0801	-0.3773**	0.1471	0.2219	-0.0072	-0.2242
R	0.6121	0.8457	0.2910	0.2645	0.6218	0.8408	0.6213	0.2689	0.6143	0.8489	0.2913	0.2652

***, ** and * significant at the 1%, 5% and 10% levels based on average t-statistics

Table 9: Results of Interaction Term (lnMV*MP) or (lnMV*FP) based on Firm Size

Firm Size	Top 20%		20%		20%		20%		Bottom 20%	
	Coefficient	<i>P-value</i>								
Group A: Monetary Policy Variables										
Monetary Base (lnMV*MB)										
TR	0.2131*	0.0844	-0.0091*	0.0848	0.0182	0.9891	0.0568**	0.0406	0.9101*	0.0925
TV	0.3997*	0.0984	0.4753***	0.0123	1.1667	0.1102	1.0689***	0.0102	2.9798***	0.0102
ILLIQ	-0.0297***	0.0119	0.0173	0.6886	0.0227	0.4962	0.0279	0.1359	-0.5940***	0.0102
TPI	0.0397***	0.0107	0.0252	0.1810	0.0197	0.8987	0.0202	0.3036	0.0131	0.6608
Bank Rate (lnMV*BR)										
TR	0.0287	0.2932	0.0223	0.6084	0.0197	0.8798	-0.0191	0.5689	-0.0088	0.0869
TV	0.0335*	0.0845	0.0332***	0.0139	-0.0153	0.0102	-0.0122***	0.0102	-0.0870***	0.0102
ILLIQ	0.0187***	0.0176	0.0199	0.9624	0.0201	0.5106	0.0202	0.2210	-0.0008***	0.0102
TPI	0.0193***	0.0103	0.0198	0.1236	0.0200	0.8834	0.0200	0.1379	0.0195	0.2817
Short-term Interest Rate (lnMV*SIR)										
TR	0.0116**	0.0463	0.0208***	0.0068	0.0197***	0.0102	0.0209	0.6716	-0.0551	0.1206
TV	0.0024	0.1074	0.0421	0.2107	0.0702	0.1542	0.0651***	0.0102	-0.1478***	0.0102
ILLIQ	0.0179***	0.0129	0.0199	0.6687	0.0201	0.5105	0.0203	0.1242	-0.0057***	0.0103
TPI	0.0208***	0.011	0.0202	0.2389	0.0199	0.9444	0.0200	0.1825	0.0198	0.8251

Group B: Fiscal Policy Variables										
Government Expenditures (lnMV*GE)										
TR	0.0146**	0.0404	0.0193*	0.0854	0.0199	0.9766	0.0202	0.4637	0.0416*	0.0892
TV	0.0100	0.1017	0.0308	0.2124	0.0477***	0.0102	0.0452***	0.0102	0.0897***	0.0102
ILLIQ	-0.0188***	0.0119	-0.0199	0.6693	-0.0201	0.4879	-0.0202	0.1404	-0.0057***	0.0103
TPI	-0.0205***	0.0107	0.0201	0.1815	0.0199	0.9065	0.0200	0.2628	0.0199	0.6936
Government Borrowing (lnMV*GB)										
TR	0.0235	0.4055	0.0204	0.8569	-0.0201	0.9709	-0.0194	0.4445	-0.0059	0.0881
TV	-0.0264***	0.1029	-0.0270***	0.0123	-0.0379***	0.0102	-0.0038***	0.0102	-0.0249***	0.0102
ILLIQ	0.0192***	0.0119	0.0199	0.6723	0.0200	0.4945	0.0201	0.1386	0.0293***	0.0103
TPI	0.0197***	0.0107	0.0201	0.1825	0.0199	0.8962	0.0200	0.2790	0.0201	0.7076
Private Borrowing (lnMV*PB)										
TR	0.0644**	0.0404	0.0096*	0.0853	0.0193	0.9875	0.0336	0.4384	0.3582*	0.0908
TV	0.1390*	0.0961	0.1913***	0.0123	0.4548***	0.0102	0.4164***	0.0102	1.1365***	0.0102
ILLIQ	-0.0012***	0.0119	-0.0189	0.6737	0.0210	0.4958	-0.0229	0.1380	-0.1960***	0.0103
TPI	-0.0275***	0.0107	-0.0219***	0.0183	0.0199	0.9034	0.0201***	0.0281	-0.0182	0.7631

Note: This table summarize the average interaction term ($Int_{i,t-1}$) found in each panel estimation for each (i)liquidity model for each market using equation 6 and 7. ***, ** and * indicates significance at the 1%, 5% and 10% levels based on average t-statistics.

Under an expansionary fiscal policy, government intervenes directly or indirectly to increase market liquidity and further to increase investors' and firms' access to credit (see Spilimbergo et al., 2009; Blanchard et al., 2010). Private borrowing, therefore, is a direct measure of the banking sector's ability and willingness to loosen credit standards following a budgetary shock (Gagnon and Gimet, 2013). On average our results from eight Asian equity markets for private borrowing are much in agreement with these empirical arguments. In our sample markets, liquidity (i.e. trading volume and turnover rate) across all portfolios increases due to a greater opportunity for private borrowing, and many of these results are statistically significant. Similarly, Amihud's (2002) illiquidity significantly decreases for larger (top 40) and smaller (bottom 40) stocks. However, results for the Bangladesh market are not as expected.

Tables 7 and 8 also report the results of other macroeconomic and control variables. Results indicate significance for factors other than monetary and fiscal policy variables on stock liquidity, for example industrial production has much influence on the (il)liquidity measures of firms. The average impact of stock returns, volatility, market capitalization, inflation and market index are all different on respective (il)liquidity ratios. However, in many instances, both liquidity and illiquidity measures are found to be strongly affected by own-lagged value, market capitalization, volatility and market index, which is consistent with the results of many earlier researchers, e.g. Chordia et al. (2005) and Fernández-Amador et al., (2013).

3.3.2 Influence of monetary and fiscal policy variables on Industries

In this final section of results, we investigate the influence of macroeconomic management policy variables on (il)liquidity at industry level. Surprisingly, a gap in our knowledge exists in this area of the relationship; however, earlier studies have highlighted the impact of monetary and fiscal policy on various characteristics of industries. For example, Ehrmann and Fratzscher (2004) show that there are strong industry-specific effects of monetary policy and the effect on market returns is likely to differ across industries. Firms in cyclical industries, capital-intensive industries, and industries that are relatively open to international trade tend to be affected more strongly. They further added that interest rate, exchange rate, and cost of capital affect the expected future earnings in different ways across industries. Dedola and Lippi (2005) find that the impact of monetary policy is stronger in industries that produce durable goods, have greater financing requirements, and lower borrowing capacity. Similarly, many recent studies, such as Belo et al. (2013) and Aghion et al. (2014) explain and document the impact of fiscal policy on different industries.

In Table 10 we present the panel estimation for (il)liquidity measures classified into three major industries traded on the Asian stock markets of our sample, namely the financial, manufacturing and service sectors. To explain the influence of monetary and fiscal policy on each (il)liquidity ratio, we report the average of median values under each category. That means we run panel regressions for four (il)liquidity measures and three monetary policy variables (i.e. base money, bank rate and short-term interest rate), calculate their median values and then take the average for all markets. Similarly, we show the average of median values of panel regressions for three fiscal policy variables (i.e. government expenditure, government borrowing and private borrowing) and four (il)liquidity measures.

Group A of Table 10 details the results related to financial industries. We see that all four monetary policy variables on average are statistically significant for three (il)liquidity measures, however, none of the fiscal variables are found to be significant. This fact is true for Bangladesh, India, Indonesia, Pakistan, Taiwan and Thailand. We find only illiquidity ratios are statistically significant for Malaysia and South Korea. Besides, individually, only the (il)liquidity of Indian, Pakistani and Taiwanese markets are affected strongly by government expenditure and private borrowing along with monetary policies. Intuitively, that indicates changes in monetary policy exert a stronger impact on the liquidity and illiquidity of the financial sector regardless of the country-specific characteristics. This is in line with our hypothesis and empirical evidence. As suggested in Blanchard et al. (2010), expansionary policy can influence the liquidity of bank and non-bank institutions. From an investment perspective, this can further increase the borrowing capacity of investors (see Spilimbergo et al., 2009; Gagnon and Gimet, 2013) as credit constraints are minimized and loose monetary policy enhances the ability of banks to generate more loans to the private sector. In addition, financial institutions are also strong investors in the equity market. Therefore, any news related to central bank policy should influence the market liquidity for financial institutions. The sensitivity of bank liquidity to the central bank's interest rate policy is also highlighted in Florackis et al. (2011). Altogether, the results reported in Group A of Table 10 are further in line with the results from the previous section. We see monetary policies have a heterogeneous impact on market liquidity and the impact largely depends on which policy is being employed. In addition, financial institutions are mostly larger firms and Table 7 shows that the liquidity of larger firms is positively (negatively) associated with expansionary (contractionary) monetary policy.

Groups B and C of Table 10 respectively summarize the average results of panel estimations of (il)liquidity measures related to the manufacturing and service sectors of our sample markets.

Results indicate the manufacturing sector is most sensitive to any policy changes by the government and central bank among all three industry categories, yet the service sector is least sensitive. Group B shows that on average fiscal policy significantly influences three (il)liquidity measures of the manufacturing sector but monetary policy only influences two. Specifically, in Bangladesh, India and Pakistan market liquidity is positively influenced by government expenditure and private borrowing; however, it is negatively affected by government borrowing. That means, in these three East Asian countries, higher government expenditure and more credit to the private sector channel into the equity market, thus increasing liquidity. Yet borrowing by the government reduces the supply of credit for private investors and thus market liquidity. As expected, the illiquidity ratios are positively related to government borrowing and negatively to government expenditure and private borrowing. This evidence is not only found in all three East Asian countries but also in Taiwan.

Among monetary policy variables, the short-term interest rate influences the turnover rate and turnover price impact of the manufacturing sector. The impact is significant in Bangladesh, India, Indonesia, Malaysia, Pakistan, South Korea and Thailand. In addition, bank rate acts as an important determinant of (il)liquidity for the manufacturing sector of Indonesia, Taiwan and Thailand. We find that these two monetary policy variables affect stock market liquidity negatively and illiquidity positively except in Pakistan, where on average the liquidity of firms is not negatively affected by bank rate or positively by money supply growth. Overall, the liquidity behaviour of the manufacturing sector in most of our sample markets is consistent with our hypothesis: liquidity rises with expansionary shocks and declines with contractionary shocks.

Earlier researchers have explained the possible linkage between monetary policy, fiscal policy and manufacturers, e.g. Ehrmann and Fratzscher (2004), Dedola and Lippi (2005) and Aghion et al. (2014). They all mention various firm specific characteristics where monetary policy has stronger influence, such as capital intensity, producing durable goods, export oriented, greater financing needs (working capital), smaller size, and lower asset tangibility. Supporting their arguments, in our sample Asian markets, the manufacturing sector mostly includes firms – pharmaceutical, garments, engineering, leather, food products - which are capital intensive, the goods are durable, they are involved in international trade and they are in need of more funds as working capital. In addition, many of these manufacturing firms are small in size, thus following the previous section they are affected significantly by policy changes. During expansionary periods investors prefer to invest in smaller company portfolios for greater price growth and we have seen the effect is even higher for fiscal policy changes (see Table 9). Finally, expansionary fiscal policy

reduces the ‘crowding out’ effect, making funds available for private firms (funds also become cheaper) and thus as asserted in Spilimbergo et al. (2009) enhances the financial health of companies.

In our sample the service sector consists mostly of IT systems, telecommunications, real estate, hospitals, and travel and leisure companies. Group C of Table 10 reports that on average Amihud’s illiquidity measures are affected by monetary policy; however, trading volume is influenced by fiscal policy. The coefficients are also statistically significant for them at different levels up to 10 percent. From the panel estimations (not reported), we see surprising results. The broad money supply positively and bank rate negatively affect Amihud’s (2002) illiquidity ratio for this sector in Bangladesh, India, Indonesia, Malaysia and Thailand. That implies contractionary (expansionary) monetary policy reduces (increases) the illiquidity for service sector stocks. This may be due to investors’ (both institutional and individual) preference for smaller stocks. As most services companies are upper middle sized firms, during expansionary monetary policy investors switch to portfolios with smaller firms (i.e. risky but profitable). In most of the eight markets we find positive responses from TPI to bank rate and short term interest rate, but when the monetary policy is measured by money supply, illiquidity is positively associated with policy changes. That means when the central bank increases the base rate and short-term interest rate, it reduces overall funds flowing into the market and the service sector become strongly affected. In addition, from the borrowers’ perspective, BR and STR directly affect the cost of capital, thus when these rates increase investors switch back to less risky investments and the liquidity of the service sector increases. This switching behaviour is also documented in Næs et al. (2011). Moreover, costs increase more for smaller firms (as they are risky) than middle sized firms. Finally, the impact of monetary policy largely depends on the liquidity measure we use and the policy we apply (see Fernández-Amador et al., 2013), hence firms and investors need to be very careful before setting up any investment strategies.

The last four columns of Group C in Table 10 illustrate the results of panel estimations of fiscal policy impact on the service sector. The results indicate that the liquidity of this sector is linked to the expansionary policy of the government and when the public has greater access to funds. For example, trading volume is negatively influenced by government borrowing and positively by private borrowing. In addition, Amihud’s (2002) illiquidity ratio is significantly affected by both government expenditure and borrowing. In effect, government expenditure reduces illiquidity and borrowing increases it for the service sector. Therefore, regardless of the firm and industry, government borrowing creates a ‘crowding out’ effect in these emerging equity markets.

Table 10: Influence of macroeconomic management variables on industries

Group A: Financial Industries

	Monetary policy variables				Fiscal policy variables			
	TR	TV	ILLIQ	TPI	TR	TV	ILLIQ	TPI
Dependent variable	0.6625***	0.7602***	0.3793***	0.4442***	0.6629***	0.7639***	0.5519***	0.5047***
Policy variable	0.1299**	0.0704***	-0.0033**	-0.0054	-0.0044	0.0097	0.0149	-3.9E-05
Return	5.7733*	-0.8814	-0.0992	-0.3283*	6.2873*	-0.4973	-0.0917	-0.0642
Standard deviation	-4.8715**	-4.7554***	-0.1212*	-1.0233***	-5.0128***	-4.7905***	0.8339*	-0.0768***
ln(Market value)	-0.0715*	0.1096	-0.0003	-0.0808	-0.0709*	0.0997	0.1116	8.87E-05
Industrial production	0.0008	0.0009	-0.0002	0.0063	0.0009	0.0016	0.0021	-4.8E-06
Inflation	-0.0066*	-0.0166***	0.0016	0.0722**	-0.0267*	-0.0300***	-0.0654	0.0007*
Stock index	0.0928	0.0227***	-0.0212***	-0.4384***	0.1194**	-0.0967***	-0.5128***	0.0023**
R-Squared	0.4949	0.7246	0.2142	0.2693	0.4984	0.6825	0.3975	0.3039

Note: This table presents the average coefficients of each variable associated with for four (il)liquidity measures of the financial sector of each equity market in our sample. The levels of significance are determined based on average t-statistics. ***, ** and * significant at the 1%, 5% and 10% levels.

Group B: manufacturing industries

	Monetary policy variables				Fiscal policy variables			
	TR	TV	ILLIQ	TPI	TR	TV	ILLIQ	TPI
Dependent variable	0.6746***	0.8326***	0.3388***	0.3105***	0.6888***	0.8268***	0.3329***	0.3018***
Policy variable	0.0271*	0.0366	0.0032	-0.0369*	0.0001	0.0146*	-0.0007*	0.0075**
Return	3.2892**	0.8622	-0.1553	0.2093***	3.5086	0.8198	-0.0189	0.0470***
Standard deviation	-2.0415***	-3.9912***	-0.0061***	1.3954***	-2.1066***	-3.9981***	-0.0061***	1.5733***
ln(Market value)	-0.0159	-0.0759	-0.0055	-0.0794	-0.0161	-0.0908	-0.0053	-0.0779
Industrial production	0.0019**	0.0107***	0.0001	-0.0118	0.0023***	0.0127***	-0.0001	-0.0126
Inflation	-0.0185***	0.0019	-0.0036**	-0.0747	-0.0231***	-0.0337	-0.0032**	-0.0177**
Stock index	0.1624***	-0.0717***	-0.0181***	-0.5815	0.1637***	-0.1787***	-0.0170***	-0.2434
R-Squared	0.5559	0.7243	0.2513	0.2339	0.5740	0.6697	0.2486	0.2293

Note: This table presents the average coefficients of each variable associated with for four (il)liquidity measures of the financial sector of each equity market in our sample. The levels of significance are determined based on average t-statistics. ***, ** and * significant at the 1%, 5% and 10% levels.

Group C: Service Industries

	Monetary policy variables				Fiscal policy variables			
	TR	TV	ILLIQ	TPI	TR	TV	ILLIQ	TPI
Dependent variable	0.7067***	0.9544***	0.4189***	0.2614***	0.7154***	0.9541***	0.4469***	0.3020***
Policy variable	-0.0272	-0.0293	0.0002*	-0.0587	0.0047	0.0193*	-0.0008	-0.0191
Return	3.8151*	3.3302	-0.0349*	-0.3834*	3.8867*	3.3386	-0.0344*	-0.2747**
Standard deviation	-2.9993**	-6.8585***	-0.0101	0.7696	-3.0564***	-7.0589***	-0.0105	0.7119
ln(Market value)	-0.0110	-0.0246	-0.0005	0.0340	-0.0116	-0.0287	-0.0005	0.0276
Industrial production	0.0027***	0.0114	-9.3E-05	-0.0002	0.0028***	0.0118	-0.0001	-0.0028
Inflation	0.0124	0.0939	0.0027**	0.0588	0.0100	-0.0017	0.0029**	0.0991***
Stock index	0.0497	-0.3819*	-0.0009*	-1.0404	0.0631*	-0.4669	0.0001*	-0.7349
R-Squared	0.5488	0.9035	0.2038	0.1347	0.5544	0.8889	0.2258	0.1509

Note: This table presents the average coefficients of each variable associated with for four (il)liquidity measures of the financial sector of each equity market in our sample. The levels of significance are determined based on average t-statistics. ***, ** and * significant at the 1%, 5% and 10% levels.

The average panel estimations detailed in Table 10 further show that other macroeconomic and control variables have significant impacts on market liquidity measures across industries in these markets, as explained in Eisfeldt (2004), Chordia et al. (2005), Söderberg (2008), and Næs et al. (2011). For example, on average financial institutions are strongly influenced by inflation and industrial production has a greater impact on the manufacturing sector. In particular, liquidity ratios of the manufacturing sector have a positive association with industrial production, yet inflation creates a mixed effect on the (il)liquidity measures of the financial sector. Interestingly, the (il)liquidity of both these industries depends on cyclical movement of the market (i.e. measured in the market index, INX_{t-1}) and in most cases, varies positively with liquidity and negatively with illiquidity. Lagged dependent variable, own return and volatility are the other factors influencing the (il)liquidity of each industry category.

4. Summary

This study sheds light on the role of monetary and fiscal policy as a potential determinant of liquidity for eight emerging stock markets of Asia. This study helps to understand the dynamic relationship between macroeconomic management policies and market liquidity from various perspectives for policy makers, investors and academics.

Our major findings are as follows: First, the daily price index of stocks has a long-run association with monetary and fiscal policy variables. Using Pedroni's (1999, 2004) and Kao's (1999) cross-sectional cointegration approach we find that the monthly prices of our sample stocks are associated with macroeconomic management variables. Second, both central bank policy and government policy Granger-cause stock market liquidity and there are reverse causalities that exist among these variables. In particular, money supply growth, short-term interest rate, government expenditure and private borrowing show significant Granger-causality on both liquidity and illiquidity ratios. Similarly, money supply growth, government borrowing and public borrowing are receiving bidirectional causality from (il)liquidity measures. Third, the signs of impulse response functions are well in line with our hypothesis for most of the Asian markets: expansionary monetary and fiscal policy increase overall market liquidity. However, the impact of some innovations is minimal and sometime markets behave imperfectly. The liquidity response in Taiwan is the most persistent, then in India, Indonesia, Malaysia and South Korea. Both Bangladesh and Thailand display some deviation from theoretical and empirical notions, yet Pakistan is least persistent to macroeconomic innovations.

Fourth, the results of variance decomposition imply that there are significant ‘crowding out’ and ‘cost of funds’ effects. Bank rate and short-term interest rate can explain a large fraction of the error variance of Amihud’s (2002) illiquidity. This is probably due to investors switching from risky to secure investments, as suggested by Næs et al. (2011). In addition, a higher interest rate means a higher cost of funds and thus banks, rather than lending money to individuals, provide more trading loans to institutions. Similarly, government borrowing creates competition for bank loans and thus private firms suffer. Fifth, the global financial crisis broadly had a negative impact on market liquidity and significantly increased market illiquidity. However, markets such as Bangladesh are mostly influenced by local factors and are less integrated into the world economy. Using non-financial shocks we find strong support for similar feedback on (il)liquidity measures.

Sixth, our panel estimations with fixed effects on liquidity also show that on average an expansionary (restrictive) monetary and fiscal policy also significantly leads to an increase (decrease) in the liquidity of individual stocks. Interestingly, we see that individual stock characteristics such as market capitalization play a role in this relationship. The liquidity of a size based portfolio shows a significant heterogeneous response to macroeconomic factors. Both large and small stocks portfolios are sensitive to changes in monetary and fiscal policy variables. Most importantly, the impact of expansionary or contractionary policy largely depends on the instruments used by the central bank or government. For example, money supply growth, government expenditure and private borrowing particularly increase the liquidity of smaller firms. On the other hand, bank rate, short-term interest rate and government borrowing decrease the liquidity of the smallest 20 percent of firms.

Seventh, when firms are classified into portfolios based on industrial categories, which has not been considered in prior studies, we find a heterogeneous impact of monetary and fiscal policy. On average, the liquidity of manufacturing firms is affected by changes of any policy variables, whereas financial institutions are only influenced by monetary policy and the service sector is least affected. Moreover, other macroeconomic variables are also found to be statistically significant when we sort the portfolios into sector rather than size. For example, financial institutions and the manufacturing sector are strongly associated with inflation and industrial production respectively and both of these sectors are sensitive to market conditions. Hence, monetary interventions of the central bank and fiscal interventions of the government should be considered as an important determinant for individual and sectoral stocks’ liquidity, which may help to explain the observed commonality in liquidity and variations in liquidity at the aggregate market level. Finally, as asserted in Chordia et al. (2005), monthly innovations in volatility and

liquidity itself explain a large fraction of the error variance in forecasting liquidity in our markets, suggesting past volatility and liquidity are the key variables in forecasting future liquidity.

Our findings are important for risk management officers and regulators. The former should care about the fact that (i)liquidity is negatively related to market volatility and to market returns. The effect is most pronounced on the largest and smallest firms and the financial sector. Moreover, the global financial crisis had a minimal effect on this market and (i)liquidity is significantly affected by the local rather global macroeconomic news in some countries. Regulators, on the other hand, may use this study as evidence that (i)liquidity spirals are driven by monetary and fiscal policy variables. The impacts are not homogenous and largely depend on the instruments used by the regulator. Therefore, they should be careful about applying their policy and consider the possible effect on the equity market while formulating those policies.

There are some interesting ways in which our results could be developed, e.g. (i) as suggested in Chordia et al. (2005) and Goyenko and Ukhov (2009), research is required on linking movements in liquidity across equity and fixed income markets. This is because the ultimate impact of any policy changes on liquidity depends on the relative attractiveness of other asset markets. (ii) This study assumes fixed effects and includes an interaction of policy variables (i.e. monetary and fiscal policy) and market capitalization of stocks in panel estimations. However, other issues of cross-sectional homogeneity of the slope parameters in the panel models need to be investigated and tested in future work.

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