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The Stochastic Discount Factor and Liquidity in Mexico and Chile

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Abstract

The stochastic discount factor differs whether it is estimated with only the most liquid stocks from the one estimated with the whole sample in some years in Mexico and Chile in the period 2006 to 2012. This is evidence that there is a liquidity premium associated with stocks in these countries, which is not permanent. The liquidity premium is more permanent in Mexico than in Chile in the period of study.

JEL Classifications: G11, G14

Keywords: Stochastic Discount Factor, Mexico, Chile, Liquidity Premium

1. Introduction

Market liquidity refers to both the time and the costs associated to the transformation of a given asset position into cash and vice versa. Most continuous-time arbitrage or equilibrium asset pricing models assume that the time and the cost required to transfer financial wealth into cash is zero. In practice, however, during financial crises (e.g., Asia 1997, Long-Term Capital Management [LTCM] 1998, and Subprime 2008) liquidity can decline precipitously and even temporarily dry out. Investors respond by aggressively bidding for the safest (i.e., most liquid) securities, which in turn raises their prices relative to less liquid securities. If liquidity for an entire financial market (i.e., systematic liquidity) evolves randomly, an asset's return that covaries more with systematic liquidity would yield a liquidity risk premium to compensate for an event in which the asset falls in price along with the ability to liquidate it. This conjecture is consistent with the evidence that systematic liquidity risk is priced in equity markets (Pastor & Stambaugh, 2003; Gibson & Mougeot, 2004). Therefore, market liquidity risk is the risk stemming from holding an asset for which potentially insufficient buyers exist at a given point in time, especially in bad times. Alternatively, the buyers can only be willing to buy the asset at a substantial discount of its fair value (Diderich, 2009, p. 94). The market liquidity risk can be characterized by the volatility of observed bid-ask spreads (Diderich, 2009, p. 201).

This work is divided as follows. This section is an introduction. Section 2 discusses the theoretical framework. Section 3 introduces the formal model. Section 4 discusses methodological issues. Section 5 discusses the analysis and results. Finally, the conclusions section follows.

2. Theoretical Framework

Brandon and Wang (2013) show that liquidity risk can explain the performance of equity hedge fund portfolios. They observed, similarly to Avramov, Kosowski, Naik, and Teo (2007; 2011), that before considering the effect of liquidity risk, hedge fund portfolios that incorporate predictability in managerial skills generate superior performance. This outperformance disappears or weakens substantially for most emerging markets, long/short hedge fund and event-driven portfolios once the liquidity risk is accounted for. Moreover, they show that the equity market-neutral and long/short hedge fund portfolios' "alphas" also entail returns for their service as liquidity providers. These results hold under various robustness tests. Brandon and Wang (2013) concentrates on liquidity risk stemming from the fact that equity hedge fund returns may covary with a market wide systematic liquidity risk factor. This is different from Getmansky, Lo, and Makarov (2004) and Aragon (2007), which focus primarily on illiquidity as a cost factor that induces serial correlation in individual hedge fund returns and that may also provide an explanation for their higher expected returns.

Sadka (2010) analyzes whether systematic liquidity risk is priced in the cross section of hedge fund expected returns. Sadka (2010) shows that the high-liquidity risk exposure hedge fund portfolio (top decile) has a statistically significant 6 percent higher annual return, on average, than the low-liquidity risk exposure hedge fund portfolio (bottom decile) during the 1994–2008 period. In contrast, Avramov *et al.* (2007; 2011) assume that an individual hedge fund return process is generated by a single equity risk factor that is, in various ways and to some degree, predictable (alpha, beta, and return). They exploit this predictability to obtain hedge fund portfolios that deliver significantly positive alphas (relative to the FH7 Fung and Hsieh (2004) benchmark). Sadka (2010) raises the possibility that the alphas in Avramov *et al.* may be systematic liquidity risk premia. On the other hand, Avramov *et al.* predict individual hedge fund alpha, beta, and return with variables unrelated to liquidity.

Cao, Chen, Liang, and Lo (2013), as in Avramov *et al.* (2007; 2011), exploits the predictability of the return process for each individual hedge fund to form optimal hedge fund portfolios that outperform the FH7 benchmark and it shows that many hedge funds exploit their ability to time (i.e., predict) liquidity to decrease (increase) their single equity factor exposure as liquidity decreases (increases). Furthermore, while Avramov *et al.* provide direct evidence of the predictability of the hedge fund return process, Cao *et al.* (2013) provide indirect evidence that hedge fund managers can predict liquidity. The top decile of liquidity-timing hedge funds has alphas as much as 9.5 percent per year above the lowest decile of liquidity timers.

Amihud and Mendelson (1986) uses Fama and MacBeth's (1973) approach to evaluate the impacts of the rate of return and risks of the market to odds ratio between selling and buying prices for a portfolio of NYSE stocks from the period 1960 to 1980. They found that increasing one percent the odds ratio between selling and buying prices increases the risk per month 0.211 percent. Moreover, they observed that the slope coefficient of the odds ratio between selling and buying prices is positive.

Chan and Faff (2005) have a similar approach as Amihud and Mendelson (1986). They consider as factors the ratio book value to market value, size of the portfolio and liquidity, similar to Fama and French (1992), for Australia stocks during the 1989-1998 period. The premium risk of the yearly turnover ratio is over 20 percent. Their findings provide strong evidence for the important role of liquidity in the Australian stock market.

Acharya and Pedersen (2005) analyze the impact of liquidity as an adjustment to the CAPM model for the NYSE and AMEX from June 1962 to 1999. They estimate that the impact of liquidity risk is 1.1 percent per year and the impact on the rate of return of liquidity is 3.5 percent per year. Thus, the overall effect of liquidity is 4.6 percent per year.

Wang and Di Iorio (2007) apply Fama and French's (1992) model. They observe the impacts of other factors in addition to a market factor, liquidity being one of them, which is measured by the turnover ratio. They analyze the Chinese stock market from 1994 to 2002. Their results show that liquidity is negative correlated with the rate of return of stocks.

Interpretation of results must be carefully analyzed. There is some contradictory evidence in the literature. For example, short term performance persistence, as the one documented in Agarwal and Naik (2000) and other papers can be simply traced to illiquidity-induced serial correlation in hedge fund returns (Getmansky *et al*., 2004), as they showed with a return-smoothing model.

3. The Model

The consumer-portfolio model, initially formulated by Merton (1969) and Samuelson (1969), is ample discussed in the literature with slightly different variations. In the problem, a consumer can trade freely in assets i and maximizes the expected value of a discounted time-separable utility function, as in Campbell, Lo, and MacKinlay (1997, p. 293),

$$Max\mathrm{E}_{\mathrm{t}}[\sum_{\mathrm{j=0}}^{\infty}\delta^{\mathrm{j}}\mathrm{U}(\mathrm{C}_{\mathrm{t+j}})], \tag{3.1}$$

where δ measures the personal time preference, C_{t+j} is the investor´s consumption in period $t+j$, and $U(C_{t+j})$ is the period utility of consumption at $t+j$. Wealth W_t at t, as in Pennacchi (2008, pp. 106), satisfies the following relation

$$W_{t+1}=\sum_{i=1}^{I}((R_{i,t}-R_{f,t})*w_{i,t}+R_{f,t})(W_t+y_t-C_t) \tag{3.2}$$

where $w_{i,t}$ is the proportion invested in risky asset i of the total wealth in period t, $R_{i,t}$ is the return of risky asset i in period t, $R_{f,t}$ is the return of the risk free asset in period t and y_t is exogenous income wealth that the individual receives at period t. Notice that in this formulation, if the rate of return is zero, the return is one, which indicates that wealth is inter-temporally conserved.

The optimal consumption and portfolio plan must satisfy that the marginal utility of consumption today is equal to the expected marginal utility benefit from investing one monetary unit in asset i at time t, selling it at time $t+1$ for $R_{i,t+1}$ and consuming the proceeds,

$$U'(C_t)=\delta E_t(R_{i,t+1}U'(C_{t+1})\big|\Psi_t), \tag{3.3}$$

where Ψ_t is the information available to the individual at time t, a subset of the information available at time t, A_t. Dividing both sides in (3.3) by $U'(C_t)$, we get

$$1=E_t(R_{i,t+1}m_{t+1}\mid\Psi_t) \tag{3.4}$$

where the stochastic discount factor m_{t+1} is equal to the stochastic inter-temporal rate of substitution $\delta U'(C_{t+1})/U'(C_t)$.

Note that if the returns of the n risky assets in the economy are the vector R_t; and $\overline{1}$ is a vector of ones, relationship (3.4) can be written as

$$\overline{1}=E\left(R_t m_t\big|\Psi_{t-1}\right) \tag{3.5}$$

where R_t has an unconditional non-singular variance-covariance matrix Σ.

An implication of this model and other inter-temporal asset pricing ones is that

$$E(R_{t+1} \mid \Psi_t] - R_t^f = \frac{Cov(R_{t+1}, m_{t+1} \mid \Psi_t)}{E(m_{t+1} \mid \Psi_t)} \tag{3.6}$$

where the return on one period riskless bond is $R_t^f = 1 / E(m_{t+1} \mid \mathrm{A}_t)$ and $R_t^f \in \Psi_t$.

A benchmark portfolio can be estimated maximizing the Sharpe Ratio S_t,

$$S_t = \frac{w_t^T (R_t - R_t^f)}{w_t^T \Sigma w_t} \tag{3.7}$$

where $R_t^m = w_t^{*T} R_t$ and w_t^* are the optimum weights.

3.1 Estimation of Euler Equation of Consumption

In equilibrium, the conditional moment condition that the stochastic discount factor m_t must satisfy conditional on previous information Ψ_{t-1} is that the expected product of any return R_t with the discount factor must be equal to one,

$$E(m_t R_t \mid \Psi_{t-1}) = 1 \tag{3.8}$$

According to Hansen and Singleton (1982) the discrete-time models of the optimization behavior of economic agents often lead to first-order conditions of the form:

$$E_t\left(h(x_t, b_o)\right) = 0 \tag{3.9}$$

where x_t is a vector of variable observed by agents at time t and b_o is a p dimensional parameter vector to be estimated. Therefore:

$$E\left(h_t\left(x_t, b_o\right)\right) = E\left(R_t m_t\right) - 1 = 0 \tag{3.10}$$

Let us construct an objective function that depends only on the available information of the agents and unknown parameters b. Let $g_0(b) = E[f(x_t; z_t; b_o)]$. According to Hansen and Singleton (1982), if the model in (3.9) is true then the method of moment estimator of the function g is:

$$\mathrm{g}_T\left(b\right) = \frac{1}{T}\sum_{t=1}^{T} f\left(x_t, z_t, b\right) \tag{3.11}$$

The value of $g_T(b)$ at $b = b0$ should be close to zero for large values of T. In this paper, we follow Hansen and Singleton (1982) and choose b to minimize the function J_t

$$J_T(b) = g_T'(b) W_T g_T(b) \tag{3.12}$$

where W_T is a symmetric, positive definite weighting matrix W_T can be estimated minimizing

$$W_T = \frac{1}{T}\sum_{t=1}^{T} (f(x_t; z_t; b) f(x_t; z_t; b)) = 0 \tag{3.13}$$

The choice of weighting matrix W_T is such that it makes g_T close to zero, taking into account possible heteroscedasticity and autocorrelation (HAC) behavior.

There are two advantages of estimating non-linear Euler equation under Generalized Method of Moments (GMM) as given in Hansen and Singleton (1982):

(a) Unlike the maximum likelihood (ML) estimator, the GMM estimator does not require the specification of the joint distribution of the observed variables.

(b) The instrument vector does not need to be economically exogenous. The only requirement is that this vector be predetermined in the period when the agent forms his expectations. Both past and present values of the variables in the model can be used as instruments. Model estimator is consistent even when the instruments are not exogenous or when the disturbances are serially correlated.

To compute W_T a consistent estimator of b_o is needed. This can be obtained by initially using $W_t = I_{r\times r}$ (identity matrix) and suboptimal choice of b in minimizing $J_t(b)$ in (3.12), we get the values of b_T . By using this value of b in (3.13) we get W_T. Again by using the new values of W_T , b_T can be obtained by minimizing equation (3.12). We repeat this process until the estimates converge. According to Hansen, Heaton, and Yaron (1996) this iterative GMM process is more efficient in small sample than a simple standard two-step procedure given by Hansen and Singleton (1982).

4. Methodology

In this study, we analyze the performance of the Mexican and Chilean Stock Markets. For each asset, arithmetic returns were estimated. In each market, we consider two market indexes as benchmark: a market index and an index built with the weights of the portfolio that maximizes the Sharpe Ratio with the previous year returns. The market index used in the Mexican Market was the Total Return Index "Índice de Rendimiento Total (IRT)" and for the Chilean Market, the Santiago Stock Exchange Index "Índice de la Bolsa de Santiago (IPSA)," both indexes adjusted inclusive for cash dividends. The annual arithmetic returns and standard deviations of these indexes in the period of study are shown in Table 1.[1]

The efficient portfolio, the one that corresponds to the equilibrium in the CAPM, can be found with the weights that maximize the Sharpe Ratio when riskless lending and borrowing is allowed if the utility function is in the hyperbolic absolute risk aversion (HARA) class or if the returns only depend on their means and their variances-covariances (Elton & Gruber, 1995 pp. 98). The M-Mexico and M-Chile indexes are the result of maximizing the Sharpe Ratio each year using the return information of the previous year. In this analysis, all stocks in each country for which there was information from the previous year were considered. From 2007 to 2012, the resulting yearly M-Mexico indexes had the following number of stocks with positive weights: 20, 19, 23, 24, 25 and 19. From 2007-2012, the resulting yearly M-Chile indexes had the following number of stock with positive weights: 49, 36, 60, 138, 68 and 51. The M-indexes can be closer to the theoretical market equilibrium return than proxy indexes, such as the IPSA for Venezuela or the IPC for Mexico.

Table 1 shows the mean and standard deviation of the returns during the whole period of study and each of the analyzed years. Notice that the expected returns of IPSA and M-Chile and the ones

[1] See Valencia-Herrera (2013) for a deeper discussion on the Generalized Method of Moments (GMM) applied to the Mexican and Chilean Stock Markets.

of IRT and M-Mexico are very similar each year; however, IPSA and IRT have more volatility than M-Chile and M-Mexico, measured by its standard deviation. In 2008, there are negative returns measured by IPSA and IRT. The same happen in 2011, when the prospects of the Mexican and Chilean economies weakened. The recovery was stronger during 2009 and 2010. The growth in 2012 was small, compared with those of 2009 and 2010. Volatility increased in 2008, then it decreased in the following two years, it increased again in 2011, and it has a slowdown in 2012, in both the IRT and the IPSA.

Table 1. Mean and standard deviation of the daily market returns in Mexico and Chile

	IRT		M-Mexico		IPSA		M-Chile	
Year	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
2007-12	1.000591	0.014539	1.00565	0.005475	1.00046	0.012046	1.000613	0.004753
2007	1.000618	0.013519	1.009858	0.000254	1.00058	0.012185	1.000568	0.002521
2008	0.999248	0.022944	1.010616	0.000488	0.99916	0.018479	1.000167	0.002902
2009	1.001659	0.017057	1.000248	0.000307	1.00169	0.010248	1.00007	0.010231
2010	1.000818	0.009068	1.000799	0.000138	1.00130	0.007358	1.000163	0.004148
2011	0.99999	0.012328	1.013017	9.03E-05	0.99944	0.013888	1.000138	0.00007
2012	1.000739	0.007107	1.000194	0.000142	1.00013	0.005965	1.00015	0.000163

Daily returns. M-Mexico and M-Chile refers to an index that maximizes the Sharpe Ratio in Mexico and Chile, respectively, based on historical previous year returns with at least 50 quotes.

Source: Own elaboration

In each market, two equations were estimated using method of moments. If equation (3.8) is estimated for each return and the return for the risk free rate is subtracted for each of the returns, the following moment conditions must be satisfied:

$$E\left(m_t\left(R_{it}\right)\right)=1 \text{ and } E\left(m_t\left(R_{it}-R_t^f\right)\right)=E\left(m_t R_{i,t}^e\right)=0 \quad (4.1)$$

where $R_{i,t}^e$ is the excess return of asset i. The risk free rate used for Chile is the one day Chilean Interbank Rate, published by the Central Bank of Chile and, for Mexico, it is the 28 days Mexican Interbank Equilibrium Interest Rate (TIIE for its Spanish initials, Tasa de Interés Interbancaria de Equilibrio), published by the Bank of Mexico. If the CAPM is satisfied, m_t can be written as $a+bR_{i,t}^e$, see, for example, Cochrane (2005, p. 152):

$$1=E\left(\left(a+bR_{m,t}^e\right)\left(R_{it}\right)\right)=aE\left(R_{m,t}^e\right)+bE\left(\mathrm{R}_{m,t}^e R_{it}\right),$$

$$E\left(m_t R_{i,t}^e\right)=E\left(\left(a+bR_{m,t}^e\right)R_{i,t}^e\right)=aE\left(R_{i,t}^e\right)+bE\left(R_{i,t}^e R_{m,t}^e\right)=0 \quad (4.2)$$

The excess market return is instrumented with the first three lags of the same variable, which are are statistically significant in a Garch model.

Liquidity refers to the time and the costs associated with the transformation of a position in an asset into cash and vice versa. The CAPM, as many asset pricing models, assumes that the cost and time required for transforming financial wealth into cash is zero. Actually, the transformation of a position in some financial assets into cash can be expensive, particularly, if the asset has a low frequency of trading. The liquidity can refer to a particular asset or fund or to the entire financial market (i.e. systematic liquidity). An asset with low liquidity will command a different return from an asset with higher liquidity to compensate for the lack of liquidity. Similarly, if the asset return covaries with systematic liquidity, it would yield a liquidity risk premium to compensate for an event in which the asset differs in price along with the ability to liquidate it. This conjecture is consistent with the evidence that systematic liquidity is priced in equity markets (Pastor & Stambaugh, 2003; Gibson & Mougeot, 2004).

Getmansky *et al*. (2004) and Aragon (2007) analyze individual asset liquidity. These authors focus primarily on illiquidity as a cost factor that induces serial correlation in individual hedge fund returns and may also provide an explanation for their higher expected returns. Sadka (2010) and Brandon and Wang (2013) study whether systematic liquidity risk is priced or not in the cross section of hedge fund expected returns. In Sadka (2010), a high-liquidity risk exposure hedge fund portfolio (top decile) has a statistically significant percent higher annual return, on average, than a low-liquidity risk exposure hedge fund portfolio (bottom decile) during 1994–2008.

Liquidity effects on returns can be considered using directly liquidity-risky proxy measures such as in Pastor and Stambaugh (2003) or Amihud (2002). This article follows an alternative approach, which considers the effect of liquidity on excess return measures (i.e. alphas and appraisal ratios), as in Agarwal and Naik (2004) and Getmansky *et al.* (2004). Each year, each stock in the market is classified as high liquidity stock or low liquidity one depending if the stock traded more than 200 days in the year or less. The effect of liquidity was considered in two ways: the effect on the constant or on the beta of the stochastic discount factor. The effect is statistically measured using a Chow test. The variable I has a value of one if the stock has more than 200 quotes in the year and zero otherwise. The moment conditions becomes

$$\begin{aligned} 1 &= a_o E\left(R^e_{m,t}\right) + a_g E\left(R^e_{m,t}\right) I + b_0 E\left(\mathrm{R}^e_{m,t}\, R_{it}\right) + b_g E\left(\mathrm{R}^e_{m,t}\, R_{it}\right) I, \\ 0 &= a_0 E\left(R^e_{i,t}\right) + a_g E\left(R^e_{i,t}\right) I + b_0 E\left(R^e_{i,t} R^e_{m,t}\right) + b_g E\left(R^e_{i,t} R^e_{m,t}\right) I \end{aligned} \tag{4.3}$$

5. Analysis and Discussion

If liquidity does not affect the stochastic discount factor, the coefficient of a_g and b_g in equation (4.3) must be equal to zero. Note that if the IRT index is employed as market index, in the two step estimation, a_g or b_g, or both are statistically different from zero in all years, except in 2009, see Table 2. In the iterated general of moments estimation, all years, either of the coefficients a_g or b_g is different from cero, see Table 3. There is a liquidity effect in the stochastic discount factor in Mexico. Similar conclusions follow if the M-Mexico index is used as market index. Only in 2011, neither coefficient, a_g or b_g, is statistically different from cero using the two step estimation, see Table 4. In the iterated GMM estimation, the results are not statistically different from zero in 2010 and 2011, see Table 5. It was not possible to test the hypothesis using iterated GMM in 2007 and 2008 because the algorithm did not converge for these years.

For Chile, the effect of liquidity is less clear. Using the IPSA index as market index, the stochastic discount market differs due to market liquidity only in 2008 and 2009 at 99 percent of significance level, and in 2007 and 2011 at 90 percent significance level using a two stage GMM procedure, as it can be seen in Table 6. In the iterated GMM estimation, the differences are sharper. The stochastic discount factor does not differ statistically with the liquidity of the assets in 2010 and 2012. However, only in 2011, it differs at the 95 percent statistical level, see Table 7.

Table 2. Chow test for Mexico with the IRT index, GMM with two steps

	a0		ag		b1		bg	
	Coef.	z	Coef.	z	Coef.	Z	Coef.	z
2007	3.58	3.34 ***	-4.11	-2.45 **	22.67	1.94 *	-37.19	-2.08 **
2008	2.33	5.2 ***	-2.00	-3.07 ***	-17.87	-4.76 ***	25.91	5.38 ***
2009	3.09	1.51	-2.89	-1.02	10.95	0.78	-14.75	-0.67
2010	1.71	3.62 ***	-1.05	-1.53	45.69	3.53 ***	-68.38	-3.94 ***
2011	2.76	5.53 ***	-2.65	-3.66 ***	-3.81	-0.64	5.18	0.59
2012	2.14	1.39	-1.69	-0.74	92.88	3.02 ***	-136.66	-3.54 ***

Coefficients a0, a1 from assets with more than 50 quotes in a year, ag, bg with more than 200 quotes in a year. ***, ** and *, statistically significant at the 99, 95 and 90 percent level.

Source: Own elaboration

Table 3. Chow test for Mexico with the IRT index, iterated GMM

	a0		ag		b1		bg	
	Coef.	z	Coef.	z	Coef.	z	Coef.	z
2007	3.58	3.34 ***	-4.11	-2.45 **	22.43	1.92 *	-36.83	-2.07 **
2008	2.24	5.13 ***	-1.88	-2.94 ***	-17.33	-4.74 ***	25.38	5.39 ***
2009	3.50	4.45 ***	-3.45	-3.27 ***	9.01	1	-11.45	-0.96
2010	1.71	3.63 ***	-1.06	-1.54	46.28	3.56 ***	-68.89	-3.97 ***
2011	2.76	5.53 ***	-2.66	-3.66 ***	-3.72	-0.63	5.04	0.58
2012	2.60	1.86 *	-2.35	-1.15	89.69	2.75 ***	-133.16	-3.27 ***

Coefficients a0, a1 from assets with more than 50 quotes in a year, ag, bg with more than 200 quotes in a year. ***, ** and *, statistically significant at the 99, 95 and 90 percent level.

Source: Own elaboration

Table 4. Chow test for Mexico with the M-Mexico index, GMM with two steps

	a0		ag		b1		bg	
	Coef.	z	Coef.	z	Coef.	z	Coef.	z
2007	1.03	0.32	-0.07	-0.01	1.03	NA	0.02	NA
2008	2.99	0.33	-2.95	-0.22	1.16	NA	0.06	NA
2009	3.10	3.77 ***	-2.89	-2.61 ***	2653.20	1.47	-3689.81	-1.5
2010	3.70	2.45 **	-3.78	-1.72 *	-3372.20	-1.62	4622.49	1.52
2011	0.98	26.14 ***	0.02	0.28	1.00	2.06 **	0.00	.
2012	2.71	4.26 ***	-2.49	-2.73 ***	10326.25	1.55	-15459.4	-1.59

Coefficients a0, a1 from assets with more than 50 quotes in a year, ag, bg with more than 200 quotes in a year. ***, ** and *, statistically significant at the 99, 95 and 90 percent level

Table 5. Chow test for Mexico with the M-Mexico index, iterated GMM

	a0		ag		b1		bg	
	Coef.	z	Coef.	z	Coef.	z	Coef.	z
2007	NA	NA	NA	NA	NA	NA	NA	NA
2008	NA	NA	NA	NA	NA	NA	NA	NA
2009	3.08	3.79 ***	-2.87	-2.61 ***	2691.93	1.5	-3740.58	-1.53
2010	3.41	2.29 **	-3.44	-1.59	-2869.51	-1.39	4031.52	1.34
2011	0.98	26.63 ***	0.02	0.28	1.00	2.1 **	0.00	.
2012	2.79	4.73 ***	-2.62	-3.11 ***	8640.72	1.53	-12815.6	-1.56

Coefficients a0, a1 from assets with more than 50 quotes in a year, ag, bg with more than 200 quotes in a year. ***, ** and *, statistically significant at the 99, 95 and 90 percent level.

Table 6. Chow test for Chile with the IPSA index, GMM with two steps

	a0		ag		b1		bg	
	Coef.	z	Coef.	z	Coef.	z	Coef.	z
2007	0.86	11.79 ***	0.20	1.96 *	-7.41	-1.07	9.42	1.02
2008	2.05	3.78 ***	-1.51	-1.96 *	-44.25	-3.96 ***	62.67	4.39 ***
2009	0.85	8 ***	0.24	1.5	52.64	3.53 ***	-84.05	-3.78 ***
2010	-4.80	-0.79	7.48	0.96	76.92	0.28	-72.04	-0.19
2011	1.50	4.28 ***	-0.68	-1.41	-14.88	-1.78 *	21.27	1.94 *
2012	1.01	107.21 ***	-0.01	-0.92	-1.92	-0.96	2.64	0.98

Coefficients a0, a1 from assets with more than 50 quotes in a year, ag, bg with more than 200 quotes in a year. ***, ** and *, statistically significant at the 99, 95 and 90 percent level.

Table 7. Chow test for Chile with the IPSA index, iterated GMM

	a0		ag		b1		bg	
	Coef.	z	Coef.	z	Coef.	z	Coef.	z
2007	-1.11	-2.05 **	2.88	3.92 ***	-49.26	-1.96 *	74.22	2.41 **
2008	2.09	3.87 ***	-1.57	-2.05 **	-43.46	-3.93 ***	61.63	4.36 ***
2009	-1.42	-2.02 **	3.69	3.52 ***	84.72	3.11 ***	-127.26	-3.57 ***
2010	-8.01	-1.15	11.61	1.3	-52.43	-0.15	107.78	0.24
2011	4.04	3.2 ***	-4.23	-2.44 **	-22.04	-1.12	35.13	1.37
2012	1.00	7345.59 ***	0.00	-1.24	0.00	-0.23	0.01	0.28

Coefficients a0, a1 from assets with more than 50 quotes in a year, ag, bg with more than 200 quotes in a year. ***, ** and *, statistically significant at the 99, 95 and 90 percent level.

Table 8. Chow test for Chile with the M-Chile index, GMM with two steps

	a0		ag		b1		Bg	
	Coef.	z	Coef.	z	Coef.	z	Coef.	z
2007	1.44	6.27 ***	-1.18	-1.95 *	84.06	0.51	-222.29	-0.51
2008	1.64	9.02 ***	-1.96	-3.6 ***	-211.50	-2.39 **	707.65	2.66 ***
2009	1.42	3.71 ***	-1.28	-1.11	-321.04	-0.4	969.20	0.4
2010	2.44	2.74 ***	-3.36	-1.62	23.03	0.4	-53.65	-0.4
2011	2.03	12.41 ***	-2.62	-6.66 ***	-2037.02	-1.66 *	5175.64	1.67 *
2012	1.18	170.67 ***	-0.45	.	0.14	NA	-0.33	NA

Coefficients a0, a1 from assets with more than 50 quotes in a year, ag, bg with more than 200 quotes in a year. ***, ** and *, statistically significant at the 99, 95 and 90 percent level.

Table 9. Chow test for Chile with the M-Chile index, iterated GMM

	a0		ag		b1		bg	
	Coef.	z	Coef.	z	Coef.	z	Coef.	z
2007	1.68	5.37 ***	-1.79	-2.19 **	448.65	1.44	-1235.10	-1.53
2008	1.65	8.93 ***	-1.99	-3.59 ***	-231.41	-2.54 **	762.92	2.79 ***
2009	1.50	3.73 ***	-1.51	-1.24	-367.67	-0.44	1111.51	0.44
2010	2.48	2.78 ***	-3.45	-1.66 *	17.96	0.31	-41.79	-0.31
2011	2.02	12.44 ***	-2.61	-6.65 ***	-1991.85	-1.63	5061.20	1.64
2012	NA	NA	NA	NA	NA	NA	NA	NA

Coefficients a0, a1 from assets with more than 50 quotes in a year, ag, bg with more than 200 quotes in a year. ***, ** and *, statistically significant at the 99, 95 and 90 percent level.

Using the M-market Index in Chile, the differences are weaker. Only the stochastic discount factor differs with liquidity at the 99 percent confidence level in 2008 and 2011 and, with a 90 percent confidence level, in 2007. See the estimates from the two stage GMM method in Table 8. The iterated method of moments results show that liquidity matters in 2008 and 2009 with a 99 percent confidence level, in 2007 with a 95 percent confidence level and in 2007 and 2010 with a 90 percent confidence level, see Table 9. The liquidity effect on the stochastic discount factor is more clearly noticed using the iterated GMM method than the two step one. Iterated GMM estimation has better small sample properties than two step estimation (Hansen *et al.*, 1996). However, large sample statistical properties of both estimators are roughly similar.

The sensibility of the stochastic discount factor to the index can differ with liquidity. A positive (negative) sign of b_g implies that more liquid stocks are less (more) discounted than more (less) liquid stock the higher the index. For Mexico, only results using the IRT index as market index show statistically significant differences with liquidity. The stochastic discount factor using the IRT Index as market index shows a higher discount for the more liquid assets in all years except 2008 and 2011. In 2007, 2010 and 2012, stocks sensibility statistically differs with liquidity at the 99 percent level. It is noticeable that the sensibility is the opposite in 2008, a year when the effects of the global liquidity crisis were stronger in the region. In this year, the excess sensibility to the index is positive and statistically significant at 99 percent level for the more liquid assets. It is also a year in which the IRT Index has an average return of less than one, see Table 1. Similar results are observed with the two stage and the iterated GMM methods, see Tables 2 and 3. Results using the M-Mexico Index as market index are inconclusive.

For Chile, if the IPSA Index is used as market index, only in 2009 the excess sensibility of the stochastic discount to the index to liquidity is statistically negative at the 99 percent level. Similar results are observed using the two step and iterated GMM methods, see Tables 6 and 7. In 2008, a year in which the average IPSA return rate was negative, the stochastic discount shows a greater sensibility to the index for the more liquid assets with a statistical significance of 99 percent. This result is observed with both, the IPSA and the M-Chile index as market indexes and using the two step and the iterated GMM methods, see Tables 6 to 9.

6. Conclusions and Recommendations

The stochastic discount factor can provide evidence of mispricing of assets, even though literature frequently discusses these issues using asset pricing models such as the CAPM or multifactor models. There is a liquidity premium factor in the Mexican and Chilean economies in some years of the period of study, 2006-2012. The liquidity premium is more frequently present if the pricing index is the IRT for Mexico or the IPSA for Chile. The liquidity premium factor is less frequently present if a portfolio with weights that maximize the Sharpe Ratio in the previous year is used as market index. The GMM method can give different inferences if a two stage estimator or an iterated one is considered.

In Mexico, the sensibility of the stochastic discount factor to the IRT as market index is smaller if the stocks are more liquid in all years except for 2008 and 2011. In 2008, in the deep of the credit crisis, the relation is statistically significant, but with an opposite sign: the sensibility of the stochastic discount factor to the IRT as market index is larger if the stocks are more liquid. For Chile, the differences given by liquidity in almost all of the years are not statistically significant. Only in 2009, the sensibility of the stochastic discount factor to the IPSA index as the market one is statistically lower for the more liquid stocks. The relation is reversed in 2008.

It is recommended a careful interpretation of the results considering a possible model misspecification, which can result from a non-linear factor model, missing factors, inadequate instruments or over-identification issues, which are left for further extensions.

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