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# Macroeconomic Stability, Financial Risks and Market Efficiency: Evidence for a Small and Open Economy

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# Macroeconomic Stability, Financial Risks and Market Efficiency : Evidence for a Small and Open Economy.*

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**ABSTRACT**

We analyze whether changes in external macroeconomic conditions are associated with changes in the dynamics of risk and changes in the degree of informativeness of the price system for the Chilean economy. By focusing our analysis in Chile, we hope to isolate the effect of external macroeconomic volatility from the effects that weak institutions or a poor rule of law might have over domestic markets performance. We find that the dynamics of local financial risks is associated with the evolution of the overall economic conditions of the country, with a strong reduction in idiosyncratic and systematic risks during periods of stable macroeconomic conditions. Since the level and composition of financial risks can affect the net benefits of acquiring and trading on private information, we conjecture that external economic conditions can also affect the informativeness of domestic stock prices. We find that the ability of domestic stock prices to reflect fundamentals information improves during times of stable external conditions. In spite of this, we provide evidence that traditional measures of informationally market efficiency in the $R^2$ literature fail to detect such changes and attempt to explain the reason behind the inability of these measures to properly proxy for stock price informativeness.

# 1 Introduction.

In the aftermath of the Asian crisis and the turbulent period of the early 2000s, the world economy experienced a period characterized by a reduced amount of risk and low volatility in financial markets. This scenario of low volatility was prevalent not only in developed economies, but was also characteristic of many emerging markets. In fact, and probably as a consequence of the stable macroeconomic conditions and good perspectives on the world economy at the time, emerging economies experienced a sharp decline in their levels of risk. For emerging markets as a whole, the EMBI global spread fell from 1200 basis points at the end of year 2001 to a minimum below 200 basis points in the middle of 2007. In the case of Chile, the country´s sovereign spread reached historical low levels, falling from 200 basis points in 2002 to an average level close to 100 basis points in 2007.

In this article, we study the dynamics of financial risks for the Chilean economy for the 10 years period between 1997 and 2007, a period characterized by changing macroeconomics conditions. We attempt to address two specific issues. First, we analyze whether changes in the stability of external economic conditions and country risks are associated with changes in stock market volatility in terms of the level and composition of risks markets participants are faced with; and second, given that the composition and level of financial risks affect traditional measures of stock price informativeness, we analyze whether the ability of prices to reflect fundamentals information changed during our sample period. We find that that there is a statistically significant reduction in systematic and idiosyncratic stock price return volatility after the year 2001. The timing of this reduction suggests that the change in the dynamics of local financial risks is associated with changes in the

overall macroeconomic conditions of the country. Regarding the informational content of stock prices, we find that the strength of the association between prices and fundamentals changes during our sample period, being much stronger during times of stable macroeconomic conditions and dimmed stock price volatility. By focusing our analysis in Chile we are able to isolate the effect of external macroeconomic volatility from the effects that weak institutions, a poor rule of law or a deficient regulatory framework might have over internal markets stability [Johnson et al., 2002].[1]

The rest of the paper is organized as follows. In section 2 we motivate our analysis in the context of the existing literature and briefly describe our hypotheses, methodology and main results. In section 3 we analyze the dynamics of systematic and idiosyncratic volatilities in the Chilean stock market for the 1997-2007 period. We also model the behavior of the *Chilean General Index of Stock Prices (IGPA)* in order to endogenously identify the timing of possible changes in the economic conditions of the Chilean economy that might have had an effect on the functioning, risks and efficiency of its local markets. In section 4 we analyze the dynamics of two common proxies for stock price synchronicity developed in the $R^2$ literature and argue that these measures might not be able to properly capture changes in the informational content of prices in an scenario in which idiosyncratic and systematic volatilities move in the same direction. We propose an alternative approach to

[1]Chile features an open capital market with a strong regulatory and supervisory institutional framework. The current situation of the country, regarded today as one of the countries with the soundest financial system of the region, is the result of the capital market reforms implemented in the early 2000s and the early economic reforms adopted during the 1980s and 1990s, which are considered as some of the most successful among emerging economies. In fact, according to *The Financial Development Report (2011),* Chile ranks 31 among 60 countries in terms of the overall degree of development of its financial system, being the second Latin American country with the highest ranking after Brazil, ranked 30. Furthermore, the country ranks 7 in terms of its financial stability, 27 in terms of its institutional environment and 23 in terms of its business environment, being the country with the highest rankings for this factors in the region.

asses the informational efficiency of stock prices based on the association of fundamental co-movements and stock prices co-movements. Section (5) provides concluding remarks.

## 2 Hypotheses Development and Literature Review.

The Chilean economy is a small and open economy with an open capital market integrated to the world economy. Despite its strong institutional framework, sound macroeconomic policies and rule of law, the country domestic markets and overall economic performance are widely affected by swings in external economic conditions. We analyze the behavior of Chilean domestic financial risks for the 1997-2007 period, a time frame that was first characterized by a period of high volatility and uncertainty in international markets, - the 1997-2001 period - followed by a time of enhanced economic conditions, reduced stock price volatilities and historically low levels of country risks - the 2001-2007 period. By decomposing total risk into systematic and idiosyncratic risks, we find that the composition and levels of financial risks experience significant changes after the year 2001, with a clearly identifiable reduction in both components of total risk after the year 2001.[2] Interestingly, the change in the dynamics of Chilean domestic risks is particularly sharp for firms‘ idiosyncratic risks. The fact that the period in which we observe a reduction in financial risks coincides with times of stable macroeconomic conditions gives raise to some interesting issues regarding the relation between asset price informativeness and macroeconomic conditions. Starting with the seminal work of Morck et al. [2000],

[2]Total risk for individual firms is proxied by the standard deviation of daily stock returns. Firm's idiosyncratic risk is proxied by the standard deviation of the residuals from standard market model regressions. Finally, firms' systematic risk is proxied by the square root of the difference between total return variance and idiosyncratic variance.

the $R^2$ literature proposes that greater firm specific return variation is associated with more informative prices and therefore, with a more *functionally efficient* market. [3] Roll [1988] argues that stock prices move together depending on the amount of firm level and market level information that they contain and shows that firm specific variation is not associated with market wide movements related to public announcements. Therefore, he argues, the variability in returns at the firm level is originated by informed investors trading. According to these arguments, and given that we find a significant reduction in idiosyncratic stock price return volatility after the year 2001, it seems to be the case that for the Chilean market stock prices are less informative during periods of economic stability.

The fact that systematic return volatility is lower during periods of economic stability is a result that easily conforms with intuition. However, and accepting that firm specific return variation is indeed a good proxy for the degree of price informativeness, why it would be then the case that prices are less informative during periods of reduced volatility in macroeconomic conditions? Information about firms fundamentals is capitalized into stock prices through the trading activities of informed traders, and these traders decide to become more or less informed depending on the net benefits of gathering and trading on private information. Grossman and Stiglitz [1980] conjecture that the more individuals are informed, the more informative is the price system, and that the number of informed individuals will basically depend on the costs of acquiring information and on how much noise there is the economy that interferes with the informative role of prices. Accordingly, a plausible explanation is that, *ceteris paribus*, during periods of calm the net benefits of

[3] Tobin (1982) proposes the *Functional Form of the* Efficient Market Hypothesis *(EMH)* developed by Fama [1970], which states that in a *functionally efficient market*, stock prices help investors to allocate their capital to those activities in which the return on their invested capital is maximized.

acquiring information are lower than during periods of high volatility and uncertainty, situation that leads less individuals to become informed and, consequently, to less informative prices. But, at the same time, during periods of economic stability there is less noise in the economy that hampers the ability of prices to convey information to investors, situation that results in more informative prices. In the stylized model developed by Grossman and Stiglitz [1980] an increase in noise -that in our case may be the result of increased levels of uncertainty in macroeconomic conditions- reduces the informativeness of prices, but it increases the returns to information and induce more market participants to become informed, which in turn results in a more informative price system. In their equilibrium, these two effects exactly offset each other, so an increase in noise turns out to have no effect on price informativeness. In the end, whether the price system becomes more or less informative as a consequence of an increase in macroeconomic volatility can be regarded as an empirical issue.

Another possible explanation to our findings that lower levels of stock price return volatility is observed during the period of stable economic conditions is based on the effects that these conditions might have on country risks, which in turn have been shown to affect return volatility. Specifically, Bartram et al. [2009] argue that there are two opposite channels through which country risk can affect the levels of risk of domestic markets; first, higher macroeconomic volatility weakens firms incentives to take on riskier activities, so a lower idiosyncratic volatility should be observed in riskier countries; or, secondly, country risks generate firm level risks that firms are not able to diversify away, which in turn induces a positive relation between country risk and idiosyncratic risk. Extending the cross sectional findings in Bartram et al. [2009] to the time series evolution of country char-

acteristics, it is possible to argue that the reduction in country risk, induced by the more favorable economic conditions prevailing during the second half of our sample period, may have affected the level and composition of financial risks in the Chilean economy. Accordingly, our results are in line with the hypothesis that external macroeconomic conditions, through their effects on country risks, generate firm level risks that firms cannot diversify away.

The conjecture that higher country risks results in higher idiosyncratic risks - by inducing firm-specific shocks that firms cannot mitigate - and the hypothesis that during periods of macroeconomic stability the net benefits of becoming informed is lower and dominates the noise effect on prices informativeness, are both consistent with our findings in the sense that both lead to the same observational consequences if the Chilean stock market effectively became less functionally efficient after 2001. The natural question that arises is whether the domestic financial market in fact became less efficient during the period of macroeconomic stability.

As a first approach, we resort to the $R^2$ literature to analyze the evolution of market efficiency in the Chilean stock market. Beginning with Morck et al. [2000], a large amount of research has evolved around the concept of stock price synchronicity measures of market efficiency.[4] Following these methodologies we provide statistical evidence that suggests that stock price synchronicity did not experience a substantial change during our sample period. If synchronicity is indeed a good measure of the informational efficiency of stock prices, the informational content of prices must have remained unchanged through

[4]The $R^2$ literature has focused both on the validity of such measures as proper proxies for price informativeness as well as on the consequences of market efficiency on firms corporate policies, particularly the investment policy, when these synchronicity measures are used as proxies for efficiency.

the 1997-2007 period. It is difficult to reconcile this result with the efficiency implications that changes in macroeconomic conditions and the level of risks might have over the ability of prices to reflect fundamentals.

In recent years the interpretation of synchronicity measures and the $R^2$ from market model regressions as good proxies for the level of price informativeness and the degree of functionally efficiency of the markets has been challenged. Kelly [2005] finds that firms with low $R^2$ exhibit higher information costs and low liquidity, result that is at odds with the hypothesis that high idiosyncratic volatility is induced by informed trading. Chan and Hameed [2006] find that analyst coverage and stock price synchronicity are positively associated in emerging markets, situation that contradicts the notion that analysts produce firm specific information. Skaife et al. [2006] present international evidence that stock price synchronicity measures are not good proxies of firm-specific information. Hou et al. [2007] develop a model in which $R^2$ is inversely related to investor's overreaction to firm-specific information and provide empirical evidence that, according to the theoretical developments of their model, there is a negative relation between $R^2$ and price momentum. Bartram et al. [2009] argue that there is no necessary relation between $R^2$ and idiosyncratic volatility. Chang and Luo [2010] provide evidence that a low $R^2$ is associated with exposure to noise trading. The results of our analysis raise one more time the question of whether synchronicity based measures are indeed good proxies for informational market efficiency.[5]

We are able to identify two factors that might explain to some extent the failure of

[5]In favor of the appropriateness of these measures, see for instance Morck et al. [2000], Durnev et al. [2003], Piotroski and Roulstone [2004], Bris et al. [2007], Saffi and Sigurdsson [2011], and Hutton et al. [2009].

synchronicity measures to properly proxy for the informational content of prices in the Chilean stock market. On the one hand, as explained by Bartram et al. [2009], because the $R^2$ from market model regressions is affected by both the idiosyncratic and the systematic components of risk, it is possible for the $R^2$ to move in the same direction as idiosyncratic risk if changes in systematic risk have the same direction and dominate the changes in idiosyncratic risk. Hence, this simultaneous reduction in both the systematic and idiosyncratic components of risk is likely to be the reason of the observed stability of this variable during our sample period and of its inability to capture changes in market efficiency during periods of varying macroeconomic conditions. On the other hand, these measures neglect the relation between stock prices and fundamentals. In this sense, statistical tests based on stock price synchronicity may be misleading in assessing market efficiency because they do not control for the nature of co-movements among prices. If stock price fundamentals tend to move together, then one should expect stock prices to move together in an efficient market. In order to shed some light about this issue we use an alternative approach to explore whether the link between prices and fundamentals actually changes during our sample period and find that the strength of this association differs between periods of low and high macroeconomic volatility. The results for the 1997-2000 period suggest that stock price co-movements are not associated with fundamentals co-movements. On the contrary, for the 2002 - 2007 period, our evidence suggests that the stock price synchronicity observed in the Chilean stock market can be explained by common movements in stock fundamentals, result that is robust to the estimation technique and the inclusion of year, industry, and liquidity effects. In any case, we do find a positive association between liquidity and the ability of stock prices to reflect fundmental information, which is in line

with the arguments in Chordia et al. [2008] in which liquidity improves market efficiency and against the hypothesis that liquidity, proxying for non informational trading, might negatively affect market efficiency (Tetlock [2006]).

In sum, our evidence suggests that for the Chilean economy, the period of stable economic conditions was also characterized by reduced idiosyncratic as well as systematic return volatilities, which constitute time series evidence for the cross sectional findings in Bartram et al. [2009] that country risks generate risks at the firm level that firms are not able to diversify away. Our results also increase the understanding of how global macroeconomic conditions can affect the informationally efficiency of the economy. In this sense, during the period of low macroeconomic volatility and risks, we find that Chilean stock prices tend to be more efficient in reflecting firms fundamentals. Since the financing decisions of firms depend upon the relative costs of debt and equity financing, it is likely that changes in the level and composition of financial risks induced by changes in country risks, constitute a direct channel through which macroeconomic conditions can affect the capital structure decisions of firms at the individual level. We think this is an important issue for further research.

# 3 Changing Economic Conditions and the Dynamics of Financial Risks.

In the years that follow the Asian crisis the world economy experienced a period characterized by low risk and low volatility in financial markets. In figure 1, panels A to C, we present the evolution of the daily rate of return for the firms included in the S&P 500

index. The time series of the return shown in panel A exhibits a noticeable reduction in the volatility starting around 2002. The monthly standard deviation of the return presented in panel B and the Chicago Board Options Exchange Market Volatility Index (*VIX*) presented in panel C documents a peak of volatility at the beginning of 2002, with a sharp decline afterwords. This scenario of low volatility was characteristic of developed as well as emerging economies. For instance, and as previously mentioned, the EMBI global spread fell from 1200 basis points at the end of year 2001 to a minimum below 200 basis points in the middle of 2007. In the case of Chile, the country sovereign spread fell from 200 basis points in 2002 to an average of 100 basis points in 2007. Similarly, the Chilean *Interbank Lending Rate* and its *Monetary Policy Rate* exhibited a noticeable reduction in their variability and levels, with the selling rates of the instruments issued by the Central Bank reaching levels well below the 4% near the end of our sample period. With respect to stock price volatility, analogous to what happened with the US market, the Chilean stock market appeared to have also experienced a reduction in its volatility. In panel D of figure 1 we present the monthly standard deviation of *IGPA* index return for the 1997-2007 period.

In this section we analyze the effects that a scenario of more stable economic conditions might have had over market volatility and the level of risks in the Chilean financial markets.

## 3.1 Evolution of Risks in the Chilean Stock Market

We follow the methodology in Bartram et al. [2009] to analyze the evolution of systematic and idiosyncratic volatilities in the Chilean stock market. We start with the whole

universe of domestic Chilean firms with stock prices information in *Economatica* for the 1997-2007 sample period. Following the standard exclusions, firms with too few market prices observations per month are dropped from the sample as well as stocks that are traded only once or whose price do not move during a given week.

Using daily returns for each of the stocks in our sample, we estimate the following equation for each quarter in our sample period:

$$r_{it} = \alpha_i + \beta_{1i} r_{mt} + \beta_{2i} \left[ r_{US,t} + e_t \right] + \varepsilon_{it} \tag{1}$$

where $r_{it}$ is the return of stock $i$ at time $t$, $r_{mt}$ is the market return at time $t$, $r_{US,t}$ is the US market return, converted to Chilean currency using the exchange rate adjustment $e_t$, and $\varepsilon_{it}$ is a random disturbance. Adopting the definitions in Bartram et al. [2009] for the various measures of risk, we use the standard deviation of daily stock returns, $\sigma_{ir}$, as a measure of total risk for individual firms. We then decompose total risk into systematic risk and idiosyncratic risk. For each firm, idiosyncratic risk $(\sigma_{i\varepsilon})$ is proxied by the standard deviation of the residual $(\varepsilon_{it})$ from market model regressions obtained from equation (1). Systematic risk $(\sigma_{is})$ is proxied by the square root of the difference between total return variance and idiosyncratic variance; that is $\sqrt{\sigma_{ir}^2 - \sigma_{ie}^2}$. We obtain aggregate measures of idiosyncratic risk and systematic risk for the market as a whole by averaging each type of risk for all our sample firms for each quarter in our sample.

We find that the median (mean) yearly idiosyncratic risk for the whole sample of firms is 37.05% (45.37%), which compares to a value of 38.7% (43.6%) for non-US firms obtained by Bartram et al. [2009]. For the annual value of systematic risk, we obtain a median (mean) of 16.35% (19.82%), compared to 18.5% (20.1%) obtained by Bartram et al.

[2009] for their sample of non-US firms. Accordingly, the levels of the aggregate systematic and idiosyncratic risks we find are in line with international risk levels documented recently in the literature.

In figure 2 we present the distribution plots for the estimated total, idiosyncratic and systematic volatilities at the firm level. Observations correspond to firm-year observations and were obtained by taking the median of the corresponding measure for each firm in each year. In panel A it is possible to observe a clear shift in the distribution of firm level total risk to smaller levels of volatility around 2001. The same situation seems to hold for idiosyncratic volatility in panel B. In panel C, for the firm level measure of daily systematic risk, there also seems to be a shift to lower levels of volatility, even though the change is not as evident as for the cases of total and idiosyncratic volatilities.

The visual inspection of figure 2 suggests that the level of financial risks that Chilean domestic firms were faced with indeed changed at the beginning of the last decade, being 2001 the year of the structural break in the dynamics of risks. We carry on formal tests to analyze the behavior of financial risk through out our sample period. We perform our analysis on both the idiosyncratic and systematic components of risks. Results are presented in table 1. As can be seen in panel A, the aggregate measures of risk exhibit a significant reduction from the pre to the post 2001 periods. The mean of the daily aggregate idiosyncratic risk falls from 2.85% to 2.52%, while its median falls from 2.5% to 2.16%. A Kolmogorov-Smirnov non-parametric test strongly rejects the null hypothesis of equality of distributions for these risk measures between the two periods. Aggregate systematic risk experienced a milder but still significant reduction between both sub sample periods, with its mean falling from 1.23% to 1.06% and its median falling from 1.13% to 0.87%.

In panel B of table 1, for the firm level measures of risk, it is possible to observe that both the idiosyncratic risk and the systematic risk experience a significant reduction after the year 2001. For the former, the mean and median values drops from 2.62% and 2.45% to 2.15% and 1.92%, respectively. For the later, the mean drops from 1.15% to 0.92%, and the median falls from 1.08% to 0.86%.[6]

We conclude that there is a statistically significant reduction in systematic and idiosyncratic risk after the year 2001 in the Chilean economy. Interestingly, the period of dimmed volatility in our sample correspond to the period of stable macroeconomic conditions. One possible explanation for this phenomenon is based on the effects that macroeconomic conditions might have on country risks. Bartram et al. [2009] provide cross sectional evidence in favor of the hypothesis that firms have limited ability to diversify away firm level risks induced by the overall riskiness of their countries, which in turn should result in a positive relation between country risk and idiosyncratic risk. Extending the cross sectional findings in Bartram et al. [2009] to the time series dimension, it is possible that the reduction in country risk may have affected the level and composition of financial risks in the Chilean economy. An alternative explanation is that during periods of stable macroeconomic conditions there is less noise in financial markets and therefore the net benefits of becoming informed are lower (Grossman and Stiglitz [1980]), so less firm specific information flows into prices, which in turn reduces firm specific volatility (Roll [1988]). Even though this later explanation appears more suitable for the reduction in idiosyncratic risk only, it raises the issue of possible changes in the degree of informativeness of the price

[6]There is also some evidence of an increase in the dispersion of idiosyncratic volatility at the firm level. The standard deviation of the firm level idiosyncratic risk experienced a 40% increase from the 1997-2000 period to the 2002-2007 period, which results significant according to a variance ratio test at any standard level of significance.

system. We explore this possibility in section 4.

## 3.2 Endogenous Structural Break Date

Even though the evolution of our measures of financial risks presented in figure 2 strongly suggests a reduction of these measures starting in 2001, we attempt to statistically identify its most likely date. For this purpose, we obtain the daily closing prices of the *General Index of Stock Prices (IGPA)* of the Santiago Stock Exchange and compute the monthly return of the index, $R_t$. This index includes all the stocks traded in the exchange and is therefore broader than our sample of firms that exclude stocks that are traded rather infrequently. We aim to capture potential changes in the overall risk conditions of the market by identifying changes in the standard deviation of the time series of the index. Following the standard procedure, we model the behavior of the monthly return as a GARCH (1,1) in mean process allowing for multiplicative heteroskedasticity[7]:

$$R_t = \alpha_0 + \alpha_1 \sigma_t^2 + \alpha_2 D_{t,\tau} + \varepsilon_t \tag{2}$$
$$\sigma_t^2 = \gamma_0 + \gamma_1 \varepsilon_{t-1}^2 + \delta_1 \sigma_{t-1}^2 + \lambda_1 D_{t,\tau}$$

where $var[\varepsilon_t] = \sigma_t^2$ and $D_{t,\tau}$ is a dummy variable that takes the value of 1 if date $t$ is greater or equal than the specified date $\tau$. We look for a structural break between December 1999 and December 2003 so as to have sufficient data points before and after that window to properly asses the significance of the dummy variable $D_{t,\tau}$ which is capturing

[7] A *Lagrange Multiplier* test strongly rejects the null of no ARCH effects in the return series of the index. The $\chi^2$ statistic attains a value of 36.687 with a p-value close to zero.

the structural break in the process of $\sigma_t^2$.

We estimate equation (2) iteratively for every possible date $\tau$ between December 31st, 1999 and December 31st, 2003.[8]. For each iteration, we compute the likelihood ratio statistic $LR = -2\left[lnL^R - lnL^U\right]$, where $L^U$ corresponds to the value of the unconstrained likelihood function evaluated at the $ML$ estimates of equation 2 and $L^R$ is the value of the restricted likelihood function obtained from the estimation of equation 2 imposing the constraint $\lambda_1 = 0$. The date $t^*$ of the structural break corresponds to the iteration in which the likelihood ratio statistic is maximized:

$$t^* = \underset{\left\{\tau\,\varepsilon\,\left[\underline{T},\overline{T}\right]\right\}}{Sup} \left(LR_\tau\right) \tag{3}$$

where $\underline{T}$ corresponds to December 31st, 1999 and $\overline{T}$ corresponds to December 31st, 2003. The supreme of the $LR$ series is obtained for April 2001.

In light of these results, we re estimate model 2 imposing a structural break in April 2001. Accordingly, the dummy variable $D_{t,\tau}$ takes the value of 0 for the all the months previous to April 31st, 2001, and a value of 1 for the rest of the months. In (unreported) regressions, we find that the estimated coefficient of the conditional variance, $\alpha_1$ , is statistically insignificant, which suggests that the conditional variance of the series is not affecting the conditional mean. Similarly, the estimated coefficient for $\alpha_2$ is insignificantly different from zero. Regarding the conditional variance specification, we find that the ARCH parameter $(\gamma_1)$ is not significant; however, the estimated GARCH parameter

[8]For the first iteration, $\tau$ = 31/12/1999, so $D_{t,\tau}$ takes the value of 0 for the all the months previous to that date and a value of 1 for the rest of the months; for the second iteration, $\tau$ = 31/01/2000, so $D_{t,\tau}$ takes the value of 0 for the all the months previous to February 1st, 1999, and a value of 1 for the rest of the months; and so on.

($\delta_1$) results highly significant, with a point estimate of 0.907 and a $z$ value of 19.46. More importantly for our arguments, the estimated coefficient of the parameter $\lambda_1$ is negative and significant. Its point estimate is -1.17 with a p-value close to zero. This result constitutes evidence that there was a significant reduction in stock price volatility for the 2001-2007 period as compared to the 1997-2000 period.[9]

We conclude that there is evidence of a structural break in the economic conditions of the Chilean economy around year 2001. The endogenously obtained date of the structural break in the volatility of the series of the return of the broad stock index *IGPA* coincides with the reported reduction in idiosyncratic and systematic risks.

# 4 Synchronicity and Stock Prices Informational Content.

In the previous section, we provide evidence of a significant reduction in stock price volatility in the Chilean economy at the beginning of the last decade. According to our results, both the idiosyncratic and the systematic financial risks exhibit a significant reduction around year 2001, which is consistent with the endogenously determined date of a

[9]Following the arguments in Allen et al. [2009], we also consider the Chilean *Interbank Lending Rate* (*ILR*) as a good proxy to detect changes in local economic conditions. We conjecture that if there is a change on the dynamics of the *ILR* series, such change should reflect changes in the conditions of the economy as a whole that might have had an effect on the level of risks of local markets. On the first place, the *ILR* is based on quotations submitted by banks through a mechanism designed to reflect the prevailing conditions in the money market and it is the aim of the monetary policy. Secondly, given that the interbank market is the most immediate source of bank liquidity, the *ILR* has a direct impact on capital allocation and risk sharing among financial institutions, effects that are transmitted to the economy as a whole. Finally, the interbank credit market is affected by the external macroeconomic conditions, particularly by those that influence on country risks, making it a suitable instrument for the identification of changes in external conditions that might affect local markets. For these reasons, we also attempt to capture a structural break in the dynamics of the *ILR* and find evidence of significant change in the conditional variance of the series in November 2001. However, this result should be interpreted with caution, because as of August 9, 2001, the Central Bank nominalized the monetary policy, which should have dimmed the variability of the series.

structural break in the volatility of the Chilean broad stock price index. In our sample, the low volatility period in the financial market was also characterized by stable macroeconomic conditions. As already mentioned, there are at least two alternative hypotheses that may explain this phenomenon, one based on the relation between the overall macroeconomic conditions of a country and the level of financial risks of its domestic markets, and the other based on the relation between the level of noise in financial markets and the benefits of acquiring firm specific information. In order to distinguish between these possible explanations, in this section we analyze the evolution of the degree of price informativeness of the Chilean economy for the 1997-2007 period. We first resort to the traditional measures developed in the $R^2$ literature to analyze the evolution of informational market efficiency. We argue against the ability of these measures to detect changes in the informativeness of prices and propose an alternative approach that explicitly takes into account the relation between prices co-movements and fundamental co-movements.

## 4.1 The Standard Synchronicity Measures in the $R^2$ literature.

In their seminal work, Morck et al. [2000] extend the arguments put forward by Roll [1988] and Grossman and Stiglitz [1980] and hypothesize that greater firm specific return variation is associated with more informative prices. These authors suggest that the degree in which stock prices move together in a given market - the *stock price synchronicity* of that market - is a good measure of its functional efficiency. They propose two measures of stock price synchronicity: first, the fraction of stocks that move together in a given time period; and second, the average $R^2$ obtained from market model regressions of the individual firms that trade on the market. The intuition for the later measure is that in a

market with a high average $R^2$, it would be difficult to distinguish between firm-specific stock price movements and market-wide price movements.

The first measure of stock price synchronicity, the fraction of stocks that move in the same direction in a given period $\tau$, is given by:

$$f_\tau = \frac{1}{T}\sum_{t=1}^{t=T}\frac{max\left[n_t^{up}, n_t^{down}\right]}{n_t^{up}+n_t^{down}} \tag{4}$$

where $n_t^{up}$and $n_t^{down}$ is the number of stocks whose prices rise and fall in period $t$, respectively, and $T$ is the number of periods in $\tau$ used to compute $f$. The second measure is based on the decomposition of the variation in individual equity returns. In line with the methodology proposed by these authors, we estimate the market model regressions (1) for all stocks in our sample. The $R^2$ of this regression for any stock $i$ measures the variation of the return of that stock explained by the market. We finally compute the quarterly weighted average $R^2$ for all available firms at time $t$ as:

$$WR_t^2 = \frac{\sum_i R_{i,t}^2 \cdot SST_{i,t}}{\sum_i SST_{i,t}} \tag{5}$$

where $SST_{i,t}$ is the is the sum of squared total variations for firm $i$ in period $t$. $WR^2$ can be interpreted as a measure of the average informational content of stock prices in the market. Morck et al. [2000] suggest that the higher the $WR^2$ of a given market, the less informative are the prices in the corresponding economy, and consequently, the less the functional efficiency of the market. We compute $f$ and $WR^2$ for the same sample of firms used in section 3 The $f$ statistic is computed on the basis of daily closing stock prices, where $t$ is weeks, $\tau$ is months and $f$ corresponds to the average fraction of stock prices

moving in the same direction during an average week, for each month, in every year of our sample period. For $WR^2$, we compute daily returns for each of our sample stocks and estimate market model regressions for each quarter in our sample period. As before, stocks that were traded only once in a given week or whose price do not move during that week are dropped from the sample.

Descriptive statistics for $f$ and $WR^2$ are presented on table 2. As shown in panel A, the mean value of $f$ is 0.69, which compares to a value of 0.67 obtained by Morck et al. [2000] for the fraction of stocks moving together in Chile in the average week of 1995. The median of this statistic is 0.68 and its standard deviation is 0.13. With respect to $WR^2$, in panel B we obtain a mean value of 0.17, which compares to a value of 0.209 obtained byMorck et al. [2000] for 1995. We next split our sample in two sub periods according to the estimated date of the structural change in domestic risk conditions and repeat the analysis for each sub sample separately. The first one corresponds to the 1997-2000 period characterized by high volatility in international and domestic markets. The other sub sample corresponds to the 2002-2007 period, a period of stable economic conditions with low volatility in financial markets. We exclude 2001, the year of the estimated structural break.[10] As reported on the table, all the descriptive statistics for both $f$ and $WR^2$ are virtually the same between sample periods, which is confirmed by the results of simple $t$ tests presented at the bottom of each table. In figure 3 we present the kernel density estimate of $f$ and $WR^2$ for the 1997-2000 period and the 2002-2007 period separately. Consistently with the results presented in table 2, the distributions of $f$ and for $WR^2$ seem to be quite similar between our sub samples. We formally test the hypothesis that the distributions

[10] In any case, our results are qualitative the same if we consider the second sub sample beginning in 2001.

of $f$ and $WR^2$ have not changed through time and present the results in table 3. We are not able to reject the null hypothesis that the distributions of $f$ and $WR^2$ are the same for both sample periods. The combined $K$-$S$ (Kolgomorov-Smirnov) statistic for $f$ exhibits a p-value close to 80% and a p-value near 1 for the $WR^2$ statistic.

These results provide evidence that suggests that stock price synchronicity did not experience a substantial change during our sample period. If synchronicity is indeed a good measure of the informational efficiency of stock prices, the informational content of prices must have remained unchanged through the 1997-2007 period. It is difficult to reconcile this result with the efficiency implications that changes in macroeconomic conditions, country risks, and financial risks are likely to have over the functioning and efficiency of financial markets. We are able to identify two factors that might explain the failure of synchronicity measures to properly proxy for market efficiency. On the first place, these measures neglect the relation between stock prices and fundamentals. In this sense, statistical tests based on stock price synchronicity may be misleading in assessing market efficiency because they do not control for the nature of co-movements among prices. If stock price fundamentals tend to move together, then one should expect stock prices to move together, which in turn would be the reason behind the relatively high and stable $f$ and $WR^2$ statistics observed during our sample period. In the next section we explore whether the link between prices and fundamentals actually changes during our sample period and find that the strength of this association differs between periods of low and high macroeconomic volatility conditions. On the second place, and specifically related to $WR^2$, the changes in the level and composition of total risk reported in section (3) can affect $R^2$ measures in manners that are not directly associated with the efficiency of the

market or the informativeness of stock prices. As explained by Bartram et al. [2009], there is no necessary relation between firm's $R^2$ and their respective idiosyncratic volatilities, because the $R^2$ from market model regressions is affected by both the idiosyncratic and the systematic components of risk. Accordingly, it is possible for the $R^2$ to move in the same direction as idiosyncratic risk if changes in systematic risk have the same direction and dominate the changes in idiosyncratic risk. Or it might remain stable if the changes in both types of risks move in the same direction off setting each other. As can be seen in table 1, at the aggregate level, the reduction in systematic risk between the 1997-2000 and 2002-2007 sub samples is 13.8%, while the corresponding reduction in idiosyncratic risk is 11.6%. At the firm level, this figures compare to a 20% reduction in systematic risk and a 18% reduction in idiosyncratic risk. This simultaneous reduction in both the systematic and idiosyncratic risk is likely to be the reason of the observed stability of the $WR^2$ during our sample period and the inability of this measure to capture the improvement in market efficiency during the period of macroeconomic stability that we report in the next section.

## 4.2 Stock Prices and Fundamental Correlations.

In this section, we analyze whether the association between stock prices and economic fundamentals remains stable during our sample period. We attempt to shed light on the issue of whether changes in the macroeconomic conditions of a country, through its effects on the level and composition of risk in its financial markets, can affect the ability of stock prices to reflect firm fundamentals. Our analysis is based on the notion that if in an efficient market stock fundamentals move together, then one should expect stock prices to move together. We therefore analyze whether stock price correlation can be explained

by fundamentals correlation, and whether the strength of this association changes over time. With this approach, we hope to overcome the problems that might arise in assessing market efficiency due to the high synchronicity of stock prices present in the Chilean stock market, while at the same time trying to stay as close as possible to the concept of stock price co-movements prevalent in the $R^2$ literature. In any case, our analysis differs from the traditional methodology proposed in this literature in that we do not consider the synchronicity of stock prices in isolation, but related to fundamentals synchronicity.

We use the same sample of firms used to compute the levels of systematic and idiosyncratic risks in section (3) We start with the whole universe of domestic Chilean firms with stock prices information in *Economatica* for the 1997-2007 sample period and perform the usual exclusions. Using weekly returns, we compute yearly return correlations for each pair of firms in our sample. We assign a unique cross section identifier to each pair of firms. We then hand collect financial information for our sample firms from the public records of the *Chilean Securities and Insurance Supervisor -SVS-.*, the government entity responsible for maintaining transparency in publicly traded markets. Following Morck et al. [2000], we use the *return on assets - ROA-* as a proxy for firm's fundamentals. This variable, defined as earnings divided by total assets, is computed for each firm-quarter observation, and then yearly *ROA* correlations are obtained for each pair of firms in our sample. For each pair of firms, firms fundamentals correlations are then matched to their corresponding firms stock price return correlations.[11] Our final sample is a panel data set comprised by pairs of firm-year observations. For each pair of firms $i$ we have yearly stock

[11] We also used the return on investments (ROI) and the return on equity (ROE) as a proxy for fundamentals. However, there are strong limitations in the availability of data and our sample gets drastically reduced.

price return correlation (*ret_corr*) and yearly *ROA* correlations (*ROA_correl*). Our goal is to measure the strength of association between price correlation and fundamentals correlations. The panel structure of our data allows us to control for unobservable, time invariant characteristics that might be affecting the return correlation between two specific firms and that it would not have been possible to control for simply using the association between stock returns and fundamentals for a cross section of firms. Furthermore, our approach is close to the concept of synchronicity used in the $R^2$ literature - we are measuring synchronicity between pairs of firms-, which allows us to shed some light on the controversy regarding the usefulness of synchronicity measures as good proxies for market efficiency.

We also classify our sample firms according to the *International Standard Industrial Classification -CIIU-*. Since it is likely that two firms in the same industry have stronger co-movements in their prices, we include in our empirical analysis (when possible according to the corresponding specification) an industry dummy variable, *IND*, that takes the value of one if both firms belong to the same industry according to the *CIIU* classification at the two digit level. Accordingly, our most general empirical specification takes the following form:

$$ret\_corr_{it} = \beta_1 + \beta_2 ROA\_correl_{it} + \beta_3 IND_i + d_t + \alpha_i + v_{it} \tag{6}$$

where $i$ indexes firm-pair observations, $t$ is time, $d_t$ correspond to year controls, and $\alpha_i$ corresponds to time invariant unobservable characteristics specific to the $i_{th}$ pair of firms. In light of the results obtained in section (3), we split our sample in two and perform the statistical analysis for each sub sample separately. Since it is between the last few months of 2001 and the beginning of 2002 when we observe that the likelihood of a structural

break in the prevailing economic conditions is maximized, in the analysis that follows, we exclude the year 2001 when splitting our sample to have a clear distinction between our sub samples when performing statistical analysis.[12] In this way, the first sub sample corresponds to the 1997-2000 period characterized by high volatility in international and domestic markets, and the second sub sample corresponds to the 2002-2007 period, a period of decreasing volatility in global markets and stable domestic macroeconomic conditions. To avoid losing too many observations due to the rather low liquidity of the Chilean stock market and our exclusions criteria, when we consider the pre 2001 and post 2001 periods separately we do not impose the condition that firms must be the same in both sample periods. For the post 2001 period, we are able to identify 38 firms for which we have enough trading information to estimate equation (6). For the pre 2001 period, we are able to identify 41 firms for which we have enough trading information to estimate it. In any case, the considered firms in the pre and post 2001 samples represent near 80% of the total capitalization of the General Price Index as for December 2000 and December 2007, respectively.

**The 1997-2000 Period**

We first analyze whether exists unobservable heterogeneity in the relation between returns correlations and fundamentals correlation for the 1997-2000 period. We estimate equation (6) by pooled OLS and (in unreported regressions) we are able to reject the null hypothesis of no time dependencies of the error term, which suggests that the regression

[12] All our results are qualitatively unchanged if we also consider the 2001 year as a post crisis year.

model should be estimated considering the panel structure of the data. [13] On table 4 we present the results of random and fixed effects estimation of equation (6). Even numbered specifications include year effects. In the random effects estimation presented in panel A, columns (1) to (4), none of the included regressors results significant. We obtain unstable estimated coefficients for the return on assets correlation, obtaining both positive and negative point estimates (but always indistinguishable from zero), depending on the inclusion of year effects. Interestingly, the fact that two firms belong to the same industry does not seem to have explanatory power. The Wald test for global significance results insignificantly different from zero when year effects are not included, but becomes significant when year effects are included, which suggests that the explanatory power of the corresponding models comes from the inclusion of the year effects only. Finally, all the Breush-Pagan LM tests strongly reject the null of zero variance of the individual effects $\alpha_i$. The fixed effects estimation results are presented in panel B, columns (1) and (2). Again, the estimated coefficient on the *ROA* correlation variable is not robust to the inclusion of year effects. It is insignificantly different from zero in column (1) when year effects are not included in the specification, and negative and significant at the 5% level when year effects are included in column (2), which is rather a counter intuitive result. In any case,

[13] We follow Wooldridge (2002) who proposes to test for serial correlation of the residuals after the estimation by pooled OLS. If there is serial correlation, then it might be the case that the error term of the regression contains a time invariant omitted factor and, consequently, pooled OLS would not be appropriate. The test for serial correlation involves a two-stage procedure; on the first stage, equation (6) is estimated by pooled OLS. On the second stage, the lagged values of the residuals from the first stage are included as regressors in the original specification. Under the null hypothesis of no time dependencies of the error term, there are no time-invariant omitted factors and the coefficient of the lag residuals should be zero. The alternative hypothesis is that the error is an auto-regressive process of order one. To test for the individual significance of the coefficient of the lagged residuals, a standard $t$ test turns out to be an appropriate statistic. The (unreported) results of this two-stages procedure reject the null hypothesis of no time dependencies of the error term at any standard level of significance.

the economic significance of the the *ROA Correl* variable in column (2) is negligible: a one standard deviation increase in this variable induces a decrease in stock price return correlation equivalent to 2.49% of the mean value of this variable for the pre 2001 period. We also present the Hausman tests for the appropriateness of the random effects estimation. In column (1), the $\chi^2$ statistic of 6.74 is significant at the 1% level, rejecting the null hypothesis of no difference between the random effects and fixed effects estimators. In column (2), when year effects are included, the Hausman test rejects the null hypothesis at the 5% level of significance. Finally, an $F$ test rejects the null of no individual effects for both specifications at any standard level of significance.

The results for the 1997-2000 period suggest that stock price co-movements are not associated with fundamentals co-movements. For all of our specifications, including the non reported pooled OLS estimation, in only one case the coefficient of the return on assets correlation turns out to be statistically significant, but it has the wrong sign and the magnitude of the coefficient is economically insignificant.

**The 2002-2007 Period**

Similarly to the results for the pre 2001 period, (in unreported regressions) we find evidence of unobservable heterogeneity in the relation between stock price returns correlations and fundamentals correlation in the pooled OLS specification for the 2002-2007 period, which suggests that our empirical specification should take into account the panel structure of the data. In panel A of table 4, columns (5) to (8), we present the results of the random effects estimation of equation (6). Even numbered specifications include year effects. We find that fundamentals correlation seem to be strongly associated with

stock price return correlations. The *ROA Correl* coefficients are stable and highly significant across the different specifications and are not affected by the inclusion of the industry dummy. Finally, and in line with the results obtained in the pooled OLS specification regarding the existence of unobservable heterogeneity, all the Breush-Pagan LM tests strongly reject the null of zero variance of the individual effects $\alpha_i$. The fixed effects estimation of equation (6) is presented on panel B of table 4, columns (3) and (4). Specification in column (4) includes year effects. Regarding the significance of the return on assets correlation, this variable remains highly significantly different from zero, suggesting that stock prices co-movements are strongly associated with fundamentals co-movements. The value of the estimated coefficients for this variable do not differ from those obtained by the random effects specification. For both specifications, a Hausman test is not able rejects the null hypothesis of random effects. Finally, an $F$ test presented below the Hausman test on the table strongly rejects the null hypothesis that all $\alpha_i = 0$.

In summary, for the 2002 - 2007 period, the evidence suggests that the stock price synchronicity observed in the Chilean stock market can be explained by common movements in stock fundamentals. This result is robust to the estimation technique and the inclusion of year and industry effects. In what follows, we show that this result is also robust to the inclusion of liquidity controls.

**Liquidity Effects**

The liquidity of the market can influence the ability of stock market prices to reflect new information. It is reasonable to expect that in more liquid markets any new piece of relevant information be incorporated more efficiently into stock prices, because trading

can be done at a lower cost and more easily, providing investors the incentives to acquire private information by increasing the net benefits of such activity. Chordia et al. [2008] argue that liquidity stimulates arbitrage activity and, as a consequence, the incorporation of private information into prices. Sadka and Scherbina [2007] show that mispricing is related to liquidity, with less liquid stocks being more likely to be overpriced. Similarly, Wurgler and Zhuravskaya [2002] and Kumar and Lee [2006] provide evidence that mispricing is more likely to occur in more liquid stocks. Also, liquidity can affect stock returns, which can affect our results since we use the correlation between stock returns as our dependent variable in equation (6). Amihud and Mendelson [1986] provide empirical evidence of a positive relation between expected returns and the bid-ask spread, a common proxy for liquidity. Jacoby et al. [2000] develop a theoretical model consistent with this evidence and provide empirical support for their model. Amihud [2002] finds a positive relation between stock returns and illiquidity over time. In a rather different perspective, Tetlock [2006] argue that liquidity might proxy for non informational trading, and consequently, negatively affect market efficiency. He shows that more liquid markets exhibit significant price anomalies than less liquid markets do.

Given the particular characteristics of the Chilean stock market, characterized by large shareholders and a strong presence of institutional investors, we analyze whether the positive and significant relationship between stock price correlations and fundamental correlations that we obtain for the 2002-2007 sub sample are robust to the inclusion of liquidity controls. As a proxy for liquidity we use the *free float* of a firm, obtained by subtracting locked in shares from the total number of shares outstanding. In other words, the free float

of a firm is the amount of shares available to market participants to trade on.[14] For each firm in our sample, the free float is computed as one minus the percentage of outstanding shares held by controllers minus the % of outstanding shares held by pension funds. Due to strong data shortcomings both in the amount and quality of ownership information, we a re not able to compute the free float for a significant number of firms for the early years of our sample. Given this situation, and because of the fact that we do not require sample firms to be the same between our sample periods, we focus our analysis in the 2002-2007 sub period. We do not consider this a serious limitation four our purposes since we are mainly interested in the post 2001 period, which is the period in which we obtain a positive and significant association between price correlations and fundamental correlations. We obtain that the average free float for this period is barely under 34% of total market capitalization, with a median value of 30% and a standard deviation of 21%. We present the evolution of the free float for our sample firms in figure 4. As can be seen from the distribution graphs, the distribution of our firms´ free float has remained stable over time. Once the free float for each firm in our sample is computed, it is not straightforward how to include this proxy for liquidity as a control in equation (6).[15] We propose the following approach: for each year, we sort our sample firms according to their average free float and assign them to terciles; then, for every year, we consider every pair of firms in our sample and assign them to one of two groups: the *low liquidity group,* that contains pair of firms

[14]The free float methodology has been adopted by many stock indexes (the S&P 500, for instance) because of its ability to capture market movements.

[15]For instance, one way to proceed could be to add the average liquidity of the corresponding pair of stocks as a control variable. The question is whether the average free float may have an impact on the correlation of prices. It can be argued that if one of the firms has a low free float, its price will have little room to reflect new information and, therefore, will tend to move slowly. Under these circumstances, the correlation with any other stock will be low, regardless of the liquidity of the other stock.

in which both of them belong to the first (low) free float quartile, and the *high liquidity group*, which contains pair of firms belonging to the third (high) free float quartile. For the low liquidity group, the mean (median) free float is 17.5% (19%), compared to a mean (median) value of 55.5% (47.5%) for the high liquidity group. We then estimate regression (6) separately for the low liquidity and high liquidity groups.

The results of our estimation are presented in table 5. In columns (1) and (2) we report the fixed effects estimation of equation (6) and in columns (3) and (4) we present analogous results for the random effects estimation. Results of a Hausman test, reported at the bottom of the table, suggest that the random effects estimation is appropriate for both the high liquidity and low liquidity sub groups estimations. We therefore focus our discussion on the results of the random effects estimation presented in columns (3) and (4). Even though the estimated coefficients for the ROA correlation variables are significant for both the low liquidity and high liquidity sub groups, we find that liquidity has a positive impact of the strength of the association between price correlations and fundamental correlations. The coefficient for the *ROA* correlation in the high liquidity group (0.043) more than doubles that of the low liquidity group (0.019). The former is significant at any standard level of significance, while the later barely loses significance at the 5% level. Furthermore, a Wald test for the equality of coefficients $\left(\chi^2 = 24.77\right)$ between the two groups results significant at any standard level of significance. Regarding the industry dummy, we find evidence that suggests that firms that belong to the same industry have stock prices that are more correlated. This effect results more important for firms in the low liquidity sub group, for which the estimated coefficient of the dummy variable is significant at the 10% level. For the high liquidity firms, the industry dummy is indistinguishable from zero.

As a general conclusion of the analysis of market efficiency of the Chilean stock market, it seems that the ability of domestic prices to reflect fundamental information improved after the year 2001, which suggests that during periods of stable economic conditions and reduced volatility, markets tend to be more informationally efficient.[16] Regarding liquidity, the relation between market efficiency and the stability of macroeconomic conditions cannot be attributed to liquidity effects. However, we do find evidence consistent with the hypothesis that for more liquid stocks relevant information is incorporated more efficiently in their stock prices. Our results suggest that, as expected, the higher the liquidity of a stock, the stronger the association between stock price correlation and fundamentals correlation. This empirical evidence is in line with the arguments in Chordia et al. [2008] in which liquidity improves market efficiency and against the hypothesis that liquidity, proxying for non informational trading, might negatively affect market efficiency. Finally, and related to the standard measures proposed in the $R^2$ literature, our results suggest that for a sample containing episodes of high volatility and unstable macroeconomic conditions, together with episodes of calm and stability in the economy, the strength of the link between prices co movements and fundamental co movements is likely to change, situation that the traditional measures of stock price synchronicity would not be able to capture. This situation might result in a lack of power of tests based on these measures to reject the null of no changes in the informational content of prices.

[16] We also checked whether fundamentals co movements changed during our sample period. We find that the fraction of firms for which their stock fundamentals move in the same direction in a given year remains stable at a 70% level through our sample period. Furthermore, (in unreported results) we cannot reject the null hypothesis that the fraction of firms for which fundamentals move together in any given year is the same between the 1997-2000 and 2002-2007 sub samples.

# 5 Summary and Conclusions

We analyze whether changes in the stability of external macroeconomic conditions and country risks are associated with changes in the level and composition of internal market risks and the degree of informationally market efficiency of the Chilean economy for the 10 years period between 1997 and 2007. According to our analysis, there is a statistically significant reduction in systematic and idiosyncratic stock price return volatility that coincides with a reduction in country risk and macroeconomic instability. We conjecture that overall economic conditions, particularly those that affect the level of risks investors face, might affect the net benefits of acquiring and trading on private information, which in turn might affect the ability of prices to reflect information about firms. We analyze the evolution of the informationally efficiency of the Chilean stock market and find that the stock price synchronicity of the domestic market has remained relatively stable over our sample period. We argue that stock price synchronicity measures are not able to detect changes in market efficiency through time and identify two factors that explain the failure of synchronicity measures to properly proxy for market efficiency. First, these measures neglect the relation between stock prices and fundamentals; second, changes in the level and composition of total risk reported can affect $R^2$ measures in manners that are not directly associated with the efficiency of the market. We propose an alternative approach to explore whether the link between prices and fundamentals actually changes during our sample period and find that during periods of stable economic conditions and reduced stock price volatility markets tend to be more informationally efficient.

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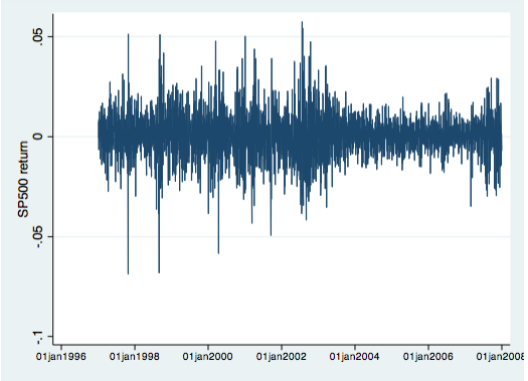
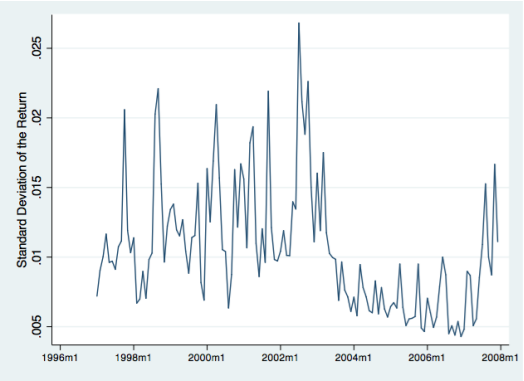
# 6 Figures and Tables

**Figure 1**. **Stock Price Volatility:** This figure presents the evolution of international volatility measures. In Panel A we present the one day rate of return of the S&P 500 index for the 1997-2007 period. In Panel B we present the monthly standard deviation of the return of the S&P 500 index for the 1997-2007 period. In Panel C the *volatility index -VIX-* is presented for the 1997-2007 period. In Panel D we present the monthly standard deviation of the return of the IGPA index for the 1997-2007 period.

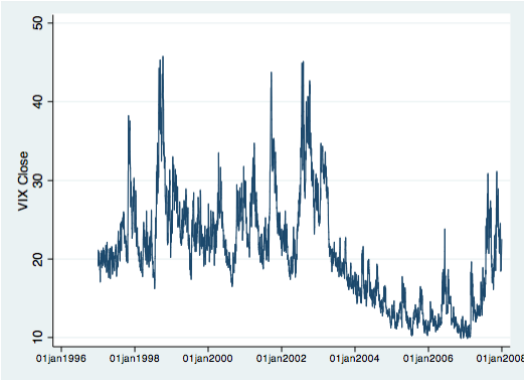
Panel A: Rate of Return S&P 500 Stock Index




Panel B: Standard Deviation S&P 500




Panel C: VIX Index




Panel D: Standard Deviation IGPA Index

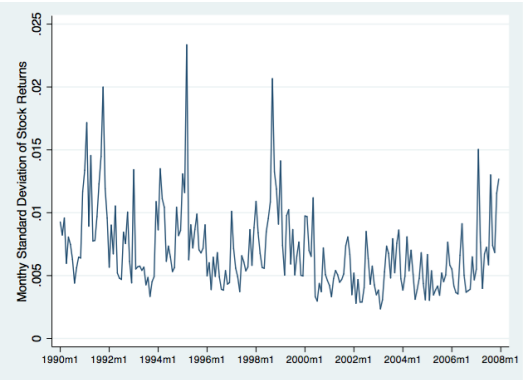
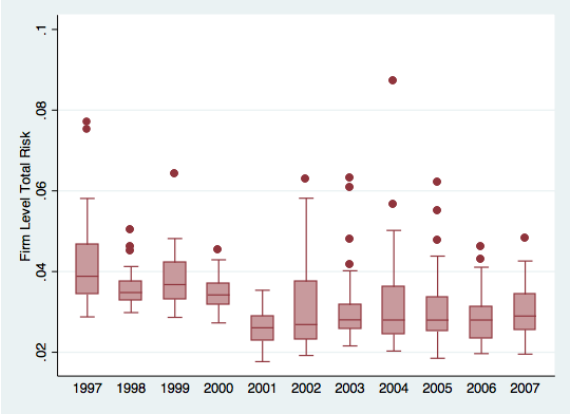
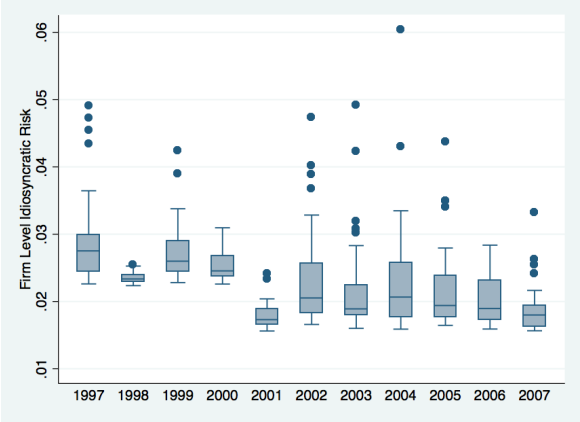


**Figure 2. Firm Level Risks:** This figure presents the evolution of firm level measures of total, idiosyncratic and systematic risk through the sample period 1997-2007. Total risk for individual firms is proxied by the standard deviation of daily stock returns. Idiosyncratic risk for each firm is the standard deviation of the residual from market model regressions obtained using daily returns. Systematic risk is proxied by the square root of the difference between total return variance and idiosyncratic variance. Observations correspond to firm-year observations obtained by taking the median of the corresponding measure for each firm in each year.

Panel A: Firm level total risk




Panel B: Firm level idiosyncratic risk




Panel C: Firm level systematic risk

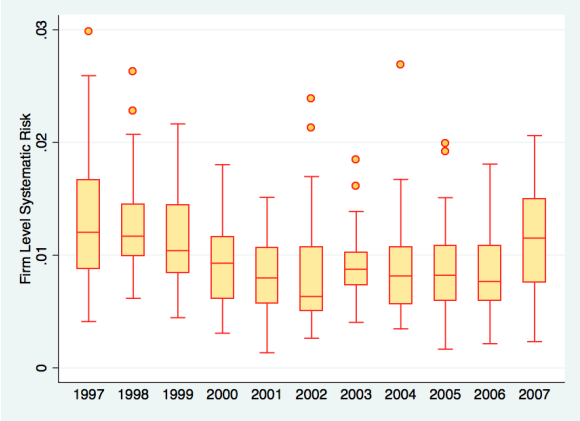
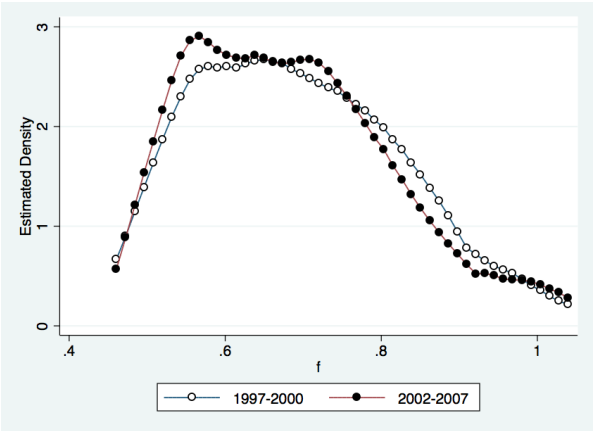


**Figure 3. Estimated Densities for $f$ and $WR^2$.** This figure presents the estimated densities for the stock prices synchronicity measures. Panel A shows the Epanechnikov kernel density estimate of $f$ for the periods 1997-2000 and 2002-2007, computed using weekly estimates of this measure. Panel B presents the Epanechnikov kernel density estimate of $WR^2$ for the periods 1997-2000 and 2002-2007 , computed using quarterly estimates of this measure.

Panel A: Kernel density estimate of $f$




Panel B: Kernel density estimate of $WR^2$

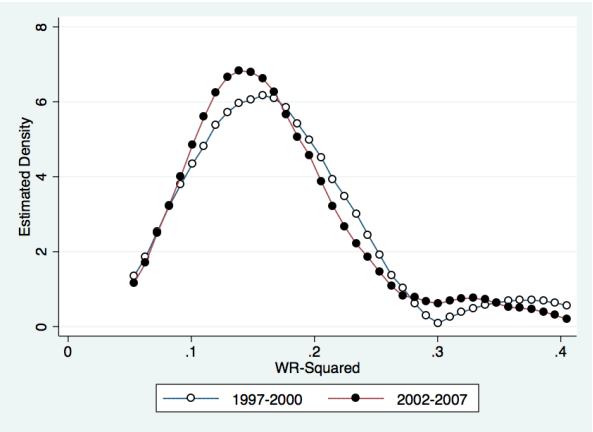


**Figure 4. Free Float:** This figure presents the evolution of the average free float of firms for the period 2002 -2007 period. The free float is computed as 1 minus the percentage of outstanding shares held by controllers minus the % of outstanding shares held by pension funds.

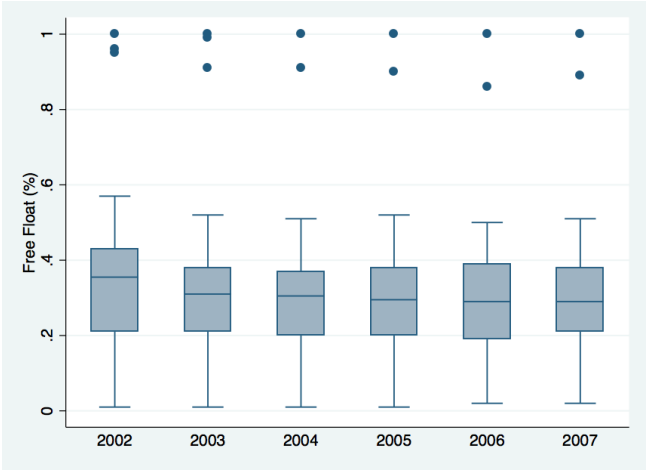


**Table 1. Descriptive Statistics and Distribution Tests for Idiosyncratic and Systematic Risk Measures:** This table presents basic descriptive statistics for idiosyncratic and systematic risk for our sample of firms during the whole sample period 1997-2007. Total risk for individual firms is proxied by the standard deviation of daily stock returns. Idiosyncratic risk for each firm is the standard deviation of the residual from market model regressions obtained using daily returns. Systematic risk is then proxied by the square root of the difference between total return variance and idiosyncratic variance. Panel A presents aggregate measures of idiosyncratic risk and systematic risk which are obtained by averaging each type of risk for all firms in each quarter. Panel B presents firm level measures of idiosyncratic risk and systematic risk. Observations in this case correspond to firm-year observations obtained by taking the median of the corresponding measure for each firm in each year.

Panel A : Aggregate Measures of Risk

| | | Aggregate Idiosyncratic Risk | | | Aggregate Systematic Risk | | |
|---|---|---|---|---|---|---|---|
| | N | Mean | Std. Dev. | Median | Mean | Std. Dev. | Median |
| Before 2001 | 14 | 2.85% | 0.0103 | 2.50% | 1.23% | 0.0049 | 1.13% |
| After 2001 | 24 | 2.52% | 0.0091 | 2.16% | 1.06% | 0.0042 | 0.87% |
| t Test for Equality of Means | | | 1.01 | | | 1.12 | |
| | | | (0.3200) | | | (0.2682) | |
| Kolmogorov-Smirnov D test | | | 0.71 | | | 0.51 | |
| | | | (0.0001) | | | (0.011) | |
| Variance Ratio Test | | | 1.2885 | | | 1.3546 | |
| | | | (0.5753) | | | (0.5074) | |

Panel B: Firm Level Measures of Risk

| | | Firm Level Idiosyncratic Risk | | | Firm Level Systematic Risk | | |
|---|---|---|---|---|---|---|---|
| | N | Mean | Std. Dev. | Median | Mean | Std. Dev. | Median |
| Before 2001 | 164 | 2.62% | 0.0046 | 2.45% | 1.15% | 0.0048 | 1.08% |
| After 2001 | 228 | 2.15% | 0.0065 | 1.92% | 0.92% | 0.0044 | 0.86% |
| t Test for Equality of Means | | | 7.91 | | | 4.84 | |
| | | | (0.0001) | | | (0.0001) | |
| Kolmogorov-Smirnov D test | | | 0.71 | | | 0.26 | |
| | | | (0.0001) | | | (0.0001) | |
| Variance Ratio Test | | | 0.4996 | | | 1.173 | |
| | | | (0.0001) | | | (0.2671) | |

**Table 2. Descriptive Statistics for $f$ and $WR^2$:** This table presents descriptive statistics for $f$ in Panel A and for $WR^2$ in Panel B. $f$ is the fraction of stocks that move in the same direction in a given period $\tau$ and is computed as $f_\tau = \frac{1}{T}\sum_{t=1}^{t=T} \frac{max\left[n_t^{up}, n_t^{down}\right]}{n_t^{up}+n_t^{down}}$ where $n_t^{up}$ and $n_t^{down}$ is the number of stocks whose prices rise and fall in period $t$, respectively, and $T$ is the number of periods in $\tau$ used to compute $f$. We compute $f$ for the sample period beginning on July 1st, 1997 and ending on December 31st, 2007, based on daily closing stock prices. In our case, $t$ is weeks, $\tau$ is months and $f$ corresponds to the average fraction of stock prices moving in the same direction during an average week, for each month, in every year of our sample period. $WR^2$ is a weighted average $R^2$ and corresponds to a measure of the average informational content of stock prices in the capital market. This statistic is computed as $WR_t^2 = \frac{\sum_i R_{i,t}^2 \cdot SST_{i,t}}{\sum_i SST_{i,t}}$, where $R_{i,t}^2$ corresponds to the multiple correlation of market model regression $r_{it} = \alpha_i + \beta_{1i} r_{mt} + \beta_{2i}\left[r_{US,t} + e_t\right] + \varepsilon_{it}$ and $SST_{i,t}$ is the is the sum of squared total variations for firm $i$ in period $t$. We compute daily returns for each of the stocks in our sample and estimate market model regressions for each quarter.

Panel A : Descriptive Statistics for $f$

| Period | N | Mean | Std. Dev. | Skewness | Kurtosis | Median | 25% Centile | 75% Centile |
|---|---|---|---|---|---|---|---|---|
| Whole Sample | 611 | 0.69 | 0.13 | 0.51 | 2.55 | 0.68 | 0.59 | 0.78 |
| Before 2001 | 221 | 0.69 | 0.13 | 0.42 | 2.42 | 0.68 | 0.59 | 0.79 |
| After 2001 | 335 | 0.68 | 0.13 | 0.60 | 2.72 | 0.67 | 0.58 | 0.77 |
| t Test for Equallity of Means | Statistic | p value (two tailed) | | | | | | |
| | 0.8526 | 0.3942 | | | | | | |

Panel B : Descriptive Statistics for $WR^2$

| Period | N | Mean | Std. Dev. | Skewness | Kurtosis | Median | 25% Centile | 75% Centile |
|---|---|---|---|---|---|---|---|---|
| Whole Sample | 42 | 0.17 | 0.07 | 1.22 | 4.86 | 0.15 | 0.12 | 0.21 |
| Before 2001 | 14 | 0.17 | 0.07 | 1.46 | 5.20 | 0.16 | 0.12 | 0.21 |
| After 2001 | 24 | 0.16 | 0.07 | 1.13 | 4.48 | 0.15 | 0.12 | 0.20 |
| t Test for Equallity of Means | Statistic | p value (two tailed) | | | | | | |
| | 0.37 | 0.7100 | | | | | | |

**Table 3. Distribution Tests for $f$ and $WR^2$:** This table presents the Kolgomorov-Smirnov distribution tests for the equality of distributions of the $f$ and $WR^2$ statistics for the periods before and after MK1. In panel A the results for the $f$ statistic are presented. In panel B the results for the $WR^2$ statistic are presented.

Panel A : Distribution Tests for $f$

| Period | H0 | Statistic | p-value |
|---|---|---|---|
| Before MK1 | The 1997 - 2000 period contains smaller values of f than the 2002-2007 period | 0.033 | 0.744 |
| After MK1 | The 1997 - 2000 period contains larger values of f than the 2002-2007 period | -0.057 | 0.422 |
| Combined K-S | The 1997 - 2000 period and the 2002-2007 period have equal distributions of f | 0.057 | 0.781 |

Panel B : Distribution Tests for $WR^2$

| Period | H0 | Statistic | p-value |
|---|---|---|---|
| Before MK1 | The 1997 - 2000 period contains smaller values of WR2 than the 2002-2007 period | 0.0536 | 0.951 |
| After MK1 | The 1997 - 2000 period contains larger values of WR2 than the 2002-2007 period | -0.125 | 0.759 |
| Combined K-S | The 1997 - 2000 period and the 2002-2007 period have equal distributions of WR2 | 0.125 | 0.999 |

**Table 4. Association between Return Correlations and Fundamental Correlations:** This table presents random effects and fixed effects estimation of the following regression equation for the pre and post MK1 periods separately: $ret\_corr_{it} = \beta_1 + \beta_2 ROA\ Correl_{it} + \beta_3 IND_i + d_t + \alpha_i + v_{it}$. Panel A shows the random effects estimation. Panel B shows the fixed effects estimation. Even numbered specifications include year effects. p-values in ( ). * Indicates significance at 10% level. **Indicates significance at 5% level. *** Indicates significance at 1% level.

Panel A: Random Effects Estimation

| | 1997-2000 Period | | | | | | | | 2002-2007 Period | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | (1) | | (2) | | (3) | | (4) | | (5) | | (6) | | (7) | | (8) |
| Constant | 0.195 | *** | 0.177 | *** | 0.195 | *** | 0.177 | *** | 0.163 | *** | 0.118 | *** | 0.159 | *** | 0.114 |
| | (0.000) | | (0.000) | | (0.000) | | (0.000) | | (0.000) | | (0.000) | | (0.000) | | (0.000) |
| ROA Correl | 0.004 | | -0.003 | | 0.004 | | -0.003 | | 0.016 | *** | 0.016 | *** | 0.017 | *** | 0.016 |
| | (0.258) | | (0.286) | | (0.260) | | (0.284) | | (0.000) | | (0.000) | | (0.000) | | (0.000) |
| Industry Dummy | | | | | 0.013 | | 0.014 | | | | | | 0.053 | *** | 0.053 |
| | | | | | (0.321) | | (0.317) | | | | | | (0.000) | | (0.000) |
| Wald $\chi^2$ | 1.28 | | 955.10 | *** | 2.26 | | 956.12 | *** | 29.06 | *** | 1649.21 | *** | 45.86 | *** | 1666.15 |
| | (0.258) | | (0.000) | | (0.322) | | (0.000) | | (0.000) | | (0.000) | | (0.000) | | (0.000) |
| Total Observations | 3280 | | 3280 | | 3280 | | 3280 | | 4218 | | 4218 | | 4218 | | 4218 |
| Number of Stock Pairs | 820 | | 820 | | 820 | | 820 | | 703 | | 703 | | 703 | | 703 |
| Breush-Pagan Test | 422.00 | *** | 798.01 | *** | 420.90 | *** | 796.52 | *** | 1536.80 | *** | 2574.90 | *** | 1481.06 | *** | 2504.21 |
| | (0.000) | | (0.000) | | (0.000) | | (0.000) | | (0.000) | | (0.000) | | (0.000) | | (0.000) |
| Year Effects | No | | Yes | | No | | Yes | | No | | Yes | | No | | Yes |

p values in ( )

**Table 4. Association between Return Correlations and Fundamental Correlations: (Con´t)**

Panel B: Fixed Effects Estimation

| | 1997-2000 Period | | | | 2002-2007 Period | | | |
|---|---|---|---|---|---|---|---|---|
| | (1) | | (2) | | (3) | | (4) | |
| Constant | 0.197 | *** | 0.180 | *** | 0.164 | *** | 0.119 | *** |
| | (0.000) | | (0.000) | | (0.000) | | (0.000) | |
| ROA Correl | 0.000 | | -0.007 | ** | 0.016 | *** | 0.015 | *** |
| | (0.979) | | (0.028) | | (0.000) | | (0.000) | |
| F | 0.000 | | 240.4 | *** | 25.290 | *** | 274.270 | *** |
| | (0.9794) | | (0.000) | | (0.000) | | (0.000) | |
| Total Observations | 3280 | | 3280 | | 4218 | | 4218 | |
| Number of Stock Pairs | 820 | | 820 | | 703 | | 703 | |
| Hausman $\chi^2$ | 6.740 | *** | 11.95 | ** | 1.130 | | 1.230 | |
| | (0.0094) | | (0.0177) | | (0.2884) | | (0.9756) | |
| F Test:$\alpha_i = 0$ | 2.670 | *** | 3.73 | *** | 4.710 | *** | 6.860 | *** |
| | (0.000) | | (0.000) | | (0.000) | | (0.000) | |
| Year Effects | No | | Yes | | No | | Yes | |

p values in ( )

**Table 5. Association between Return Correlations and Fundamental Correlations controlling for Liquidity:** This table presents random effects and fixed effects estimation of the following regression equation for the 2002-2007 period: $ret\_corr_{it} = \beta_1 + \beta_2 ROA\ Correl_{it} + \beta_3 IND_i + d_t + \alpha_i + v_{it}$. Estimation is made separately for *low liquidity* and *high liquidity* subgroups separately. Liquidity is proxied by the free float of the firms. The free float for each firm and for every year of the 2002-2007 period is computed as one minus the percentage of outstanding shares held by controllers minus the % of outstanding shares held by pension funds. For each year, firms are sorted according to their average free float and assigned to terciles. Then, for every year, every pair of firms in the sample is assigned to one of two groups: the *low liquidity group,* which contains pair of firms in which both of them belong to the first (low) free float quartile, and the *high liquidity group*, which contains pair of firms belonging to the third (high) free float quartile.

| | Fixed Effects | | | | Random Effects | | | |
|---|---|---|---|---|---|---|---|---|
| | (1) | | (2) | | (3) | | (4) | |
| | Low Liquidity | | High Liquidity | | Low Liquidity | | High Liquidity | |
| Constant | 0.162 | *** | 0.082 | *** | 0.149 | *** | 0.058 | ** |
| | (0.000) | | (0.003) | | (0.000) | | (0.040) | |
| ROA Correl | 0.018 | * | 0.028 | | 0.019 | * | 0.043 | *** |
| | (0.076) | | (0.143) | | (0.056) | | (0.000) | |
| Industry Dummy | | | | | 0.067 | * | 0.041 | |
| | | | | | (0.082) | | 0.192 | |
| F | 44.10 | *** | 8.40 | | | | | |
| | (0.000) | | (0.000) | *** | | | | |
| Wald Chi2 | | | | | 282.800 | *** | 58.440 | *** |
| | | | | | (0.000) | | (0.000) | |
| Observations | 432 | | 272 | | 432 | | 272 | |
| Number of Stock Pairs | 116 | | 67 | | 116 | | 67 | |
| Year Effects | Yes | | Yes | | Yes | | Yes | |
| Test for Equality of Coefficients | | | | | | | | |
| Wald Chi2 | 26.77 | | | | 24.77 | | | |
| | (0.000) | | | | (0.000) | | | |
| Hausman Chi2 | 4.73 | | 2.83 | | | | | |
| | (0.579) | | (0.830) | | | | | |

p values in ( )