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ABSTRACT

Creditor Protection and Stock Price Volatility*

This paper analyzes the effect of creditor protection on the volatility of stock market returns. Our application of the Tobin's q model predicts that credit protection reduces the probability of oscillations between binding and nonbinding states of the credit constraint, which result from liquidity crises and their aftermath. In this way creditor protection regulation reduces the stock market price volatility. We test this prediction by using cross-country panel regressions of the stock return volatility, in 40 countries, over the period from 1984 to 2004. Estimated probabilities of big shocks to liquidity are used as a forecast of a switch from a credit–unconstrained to a credit-constrained regime. We find support for the hypothesis that creditor protection institutions reduce the probability of oscillations between binding and nonbinding states of the credit constraint and thereby help reduce the asset price volatility.

JEL Classification: E44
Keywords: collateral, credit constrained regimes and probability of liquidity crisis

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1 Introduction

Recent literature on law and finance has emphasized the role of strong institutions, such as creditor protection, in fostering the development of financial markets. Creditor rights regulation helps mitigate the problems of information asymmetry and moral hazard between creditors and borrowers.

A central problem is that lenders are reluctant to make loans because they cannot easily determine whether a prospective borrower has resources to repay the loan. If the loan is made, the lender is concerned whether the borrower will engage in risky behavior that could lower the probability that the loan will be repaid. Collateral reduces this information asymmetry problem because good collateral (that is, assets that are easily valued and easy to take control of) significantly decreases the losses to the lender if the borrower defaults on the loan. Good collateral also reduces the moral hazard problem because the borrower is reluctant to engage in excessively risky behavior since he or she has now something to lose. Creditor protection enhances the ability of the lender to take control of the collateral in case of default and thereby alleviate credit constraints. Thus, creditor rights regulation helps mitigate the problems of information asymmetry and moral hazard between creditors and borrowers.

Hence, creditor right protection is shown to affect the credit cycle and credit market breadth. For example, La Porta et al. (1997) find that countries with poor creditor protection have smaller debt markets. Their findings are confirmed by Levine (2004) as well as Djankov, McLiesh, and Shleifer (2006), with broader country coverage. Burger and Warnock (2006) further find that countries with strong creditor rights have more developed local bond markets and rely less on foreign–currency bonds. Moreover, Galindo and Micco (2005) report that strong creditor rights can reduce the volatility of the credit market. Besides the impact on the macroeconomy, creditor protection also affects firms' investment and operation. Specifically, it lowers firms' borrowing costs and increases firms' value (e.g., La Porta et al. (2000) and Bae and Goyal (2003)) and reduces cash–flow risk, operating income variability, and operating leverage (e.g., Claessens, Djankov, and

Nenova (2001)).

So far, these studies have focused mainly on the credit market and little on the stock market. In this paper, we try to fill this gap by looking at how creditor rights affect the stock return volatility for market aggregates.[1] We find an empirical regularity that better creditor protection is associated with lower stock price volatility, and we attempt to investigate the mechanism behind it.

Credit protection may relax credit constraints and affect stock price volatility in two conflicting ways. On the one hand, if the credit constraint switches between binding and nonbinding regimes, better creditor protection could decrease price volatility. The intuition is as follows: Firms need to provide collateral to creditors and thus face credit constraints. When the credit constraint oscillates between binding and nonbinding due to external shocks, firms' investment, operation, and capital return will fluctuate as well. Because the probability of this oscillation can be reduced with better creditor protection, firm value (first moment) will rise, and the volatility (second moment) of firm value will decline. On the other hand, if the credit constraint always binds, then better creditor protection will increase rather than decrease price volatility. The intuition is that credit constraint would reduce the upside potential of good shocks by constraining firm growth. Creditor protection relaxes this constraint (although the constraint remains binding), and thus allows good shocks to affect a firm's operation, raising both the firm's value and the volatility.

We illustrate this double–effect argument with a standard Tobin's q investment model. We start with the free market case and derive the closed-form solution of Tobin's q (our theoretical counterpart of stock price) and its volatility. We then introduce the credit constraint, which depends on the degree of creditor protection. We show that, given a distribution of productivity shocks, weaker creditor protection causes more variation of Tobin's q, due to the credit constraint oscillation between binding and nonbinding. But the same level of creditor protection results in less variability

[1]Some studies have examined how corporate control affects the dispersion of stock prices within a market. For example, Morck, Yeung, and Yu (2000) look at the stock price co-movement within a country. They find that co-movement is more pronounced in poor economies than in rich economies, which they contribute to cross-country differences in property rights. Our work is not concerned with the idiosyncratic dispersion of stock prices, but rather with the instability in the aggregate.

of the Tobin's q under the credit–constrained regime. The prediction of the model which we take to the data is that better creditor protection reduces the effect of the changes in the probability of a liquidity shock on stock price volatility.

In the empirical part we look at aggregated stock return volatility in 40 countries over the years 1984-2004. We use a two-stage cross-country analysis to further examine the relationship between stock market volatility and creditor protection. In the first stage, we look at how creditor protection affects the probability of a liquidity crisis, defined as a large rise of the real interest rate, which serves as a proxy for the probability that a credit constraint is binding. We find that better creditor protection reduces the probability of a liquidity crisis. In the second stage, we examine whether the predicted crisis probability has an expected effect on the stock market volatility. We find that a higher crisis probability is indeed connected with larger stock return volatility. Next, we include the interaction term between the predicted crisis probability and the indicator of strong creditor rights protection and find that having strong creditor rights protection mitigates the impact of the predicted crisis probability on stock return volatility, as our model predicts.

The remainder of the paper proceeds as follows. Section 2 presents the theory. Section 3 describes the data, empirical regularity, empirical approach, and results. Section 4 summarizes the conclusions.

2 Theory

This section develops a Tobin's q investment model, which yields relationships between a creditor protection parameter, aggregate liquidity shocks, and the equilibrium stock market price volatility.

2.1 A Tobin's q Model of Stock Market Price Volatility: An Unconstrained Regime

We first consider a small open economy producing a single aggregate tradable good. The production function is Cobb-Douglas:[2]

$$Y_t = A_t K_t^{1-\rho}, \tag{1}$$

where A_t, $1-\rho$, and K_t denote an idiosyncratic productivity parameter, the distributive share of capital, and the capital stock, respectively. We assume that productivity levels follow a first-order autoregressive stochastic process:

$$\ln(A_{t+1}) = \gamma \ln(A_t) + \varepsilon_{t+1}, \tag{2}$$

where ε_{t+1} follows a uniform distribution over the region $[-1, 1]$. We denote by lowercase letters the log of variables in uppercase letters. Accordingly, equation (2) is rewritten as

$$a_{t+1} = \gamma a_t + \varepsilon_{t+1}. \tag{3}$$

A cost-of-adjustment investment technology for gross investment (Z_t) is specified by

$$Z_t = I_t \left(1 + \frac{1}{2}\frac{1}{v}\frac{I_t}{K_t}\right), \tag{4}$$

where $I_t = K_{t+1} - K_t$, denotes net capital formation (assuming zero depreciation) and $\frac{1}{v}$ is a cost-of-adjustment coefficient. As usual, gross investment exceeds net capital formation because of the additional reorganization and retraining costs that are typically associated with the installation of new capital equipment.

Producers maximize the expected value of the discounted sum of profits, subject to the available

[2]The model is based on Krugman (1998) and Frenkel and Razin (1996, Chapter 7).

production technology and cost-of-adjustment investment technology. The Lagrangian is

$$L = E\left[\Sigma_{t=0}^{\infty}\frac{1}{(1+r)^{t}}\left(A_t K_t^{1-\rho} - Z_t + q_t\left(K_t + I_t - K_{t+1}\right)\right)\right]. \tag{5}$$

We denote by r the world interest rate. The Lagrangian q_t is interpreted as the Tobin's q market value of a unit of new capital.

A first-order condition (derived with respect to I_t) is

$$1 + \frac{1}{v}\frac{I_t}{K_t} = q_t. \tag{6}$$

Linearizing $\ln\left(v\left(q_t - 1\right) + 1\right)$ yields

$$k_{t+1} = k_t + v\left(q_t - 1\right). \tag{7}$$

Another first-order condition (derived with respect to K_{t+1}) is

$$q_t = \frac{1}{1+r}\left(E_t\left[R_{t+1}\right] - \frac{1}{2}\frac{1}{v}\left(\frac{I_{t+1}}{K_{t+1}}\right)^2 + E_t\left[q_{t+1}\right]\right), \tag{8}$$

where R_{t+1} denotes the $t+1$ capital rental rate. The maximizing value rule in equation (8) implies that the cost of investing an additional unit of capital in the current period must be equal to the expected present value of the next period's marginal productivity of capital, plus the next period's induced fall in the adjustment cost of investment (resulting from the enlarged stock of capital due to current investment), plus the continuation value in the capital remaining for the entire future.

Note that from equation (1) and perfect competition in the capital market we have

$$R_{t+1} = (1-\rho)\, A_{t+1} K_{t+1}^{-\rho}. \tag{9}$$

Linearizing $\ln(R_{t+1})$, and using the expression $\pi \equiv 1 + \ln(1-\rho)$, yields

$$R_{t+1} = \pi - \rho k_{t+1} + a_{t+1}. \tag{10}$$

According to equation (6),

$$\frac{1}{v}\left(\frac{I_{t+1}}{K_{t+1}}\right)^2 = v\left(q_{t+1}-1\right)^2. \tag{11}$$

Hence, equation (8) becomes

$$q_t = \frac{1}{1+r} E_t\left((1+R_{t+1}) - \frac{1}{2}v\left(q_{t+1}-1\right)^2 + \left(q_{t+1}-1\right)\right), \tag{12}$$

At the deterministic steady state, $I_t = 0$ and $q_t = 1$. Therefore, the term $(q_{t+1}-1)^2$ around the steady state is an order of magnitude smaller than the term $(q_{t+1}-1)$. Accordingly we drop the $(q_{t+1}-1)^2$ term from equation (12) to obtain

$$(1+r)q_t = E_t\left[R_{t+1}\right] + E_t\left[q_{t+1}\right]. \tag{13}$$

Combining equations (7), (10), and (13), we get

$$q_t = \frac{(\pi + \rho v - \rho k_t + \gamma a_t \ + E_t q_{t+1})}{1+r+\rho v}. \tag{14}$$

We then solve for the equilibrium q_t by "guessing" a linear equilibrium relationship as follows:

$$q_t = B_0 + B_1 a_t + B_2 k_t. \tag{15}$$

From equations (7) and (15), we get

$$E_t q_{t+1} = B_0 + B_1\left(\gamma a_t\right) + B_2\left(k_t + v\left(q_t - 1\right)\right). \tag{16}$$

Substituting equations (15) and (16) into equation (14), we solve for B_0, B_1, and B_2 by comparing coefficients for a_t and k_t:

$$\begin{aligned} B_0 &= \tfrac{-\pi-\rho v+vB_2}{-r-\rho v+vB_2} \\ B_1 &= \tfrac{\gamma}{1+r+\rho v-vB_2-\gamma} \\ B_2 &= \tfrac{r+\rho v-\sqrt{(r+\rho v)^2+4\rho v}}{2v}. \end{aligned} \tag{17}$$

Note that the equilibrium value, q_t, is negatively related to the capital stock, k_t.

Based on equations (6) and (17), the equilibrium investment level is

$$I_{t0} = vK_t\left(B_0 + B_1a_t + B_2k_t - 1\right). \tag{18}$$

As expected, investment rises with a positive productivity shock (B_1 is positive).

2.2 A Tobin's q Model of Stock Market Price Volatility: A Credit Constrained Regime

We now analyze a Tobin's q model in the presence of credit constraints.[3]

Assume that the credit constraint is

$$I_t \ \leq \omega K_t - W_t, \tag{19}$$

where ω is a creditor protection parameter (better credit protection is associated with a larger ω).[4] The variable W_t stands for debt repayment, triggered by an exogenous aggregate liquidity shock. A firm may be required to pay back more of its existing debt when such a shock occurs. Because

[3] See the related literature of Bernanke and Gertler (1989), Hart and Moore (1994), Kiyotaki and Moore (1997), and Mendoza (2006a,b).

[4] In the literature on credit constraint and financial accelerator, the constraint tends to be based on a firm's market value $\omega q_t K_t$. However, if both q_t and K_t are endogenous as in Mendoza (2006b), then no tractable solution is available for q_t. By using ωK_t rather than $\omega q_t K_t$, we are able to provide tractable closed-form solutions for q_t and its volatility.

an aggregate liquidity shock increases W_t, it results in smaller investment, I_t.

Let us start with the constrained regime, where the investment constraint always binds, i.e.,

$$I_t \; = \omega K_t - W_t. \tag{20}$$

To simplify the analysis we assume that the aggregate liquidity shock, W_t , is permanent. That is, if the shock occurs in period t, then the credit constraint will be binding forever.

The Lagrangian with the constrained regime is

$$L' = E\left[\Sigma_{t=0}^{\infty}\frac{1}{(1+r)^t}\left(A_t K_t^{1-\rho} - Z_t + q_t\left(K_t + I_t - K_{t+1}\right) + \lambda_t\left(I_t \; - \omega K_t + W_t\right)\right)\right]. \tag{21}$$

The Lagrangian coefficient λ_t is associated with the credit constraint.

The first-order condition (derived with respect to I_t) is

$$1 + \frac{1}{v}\frac{I_t}{K_t} = q_t + \lambda_t. \tag{22}$$

Linearizing $\ln\left(v\left(q_t + \lambda_t - 1\right) + 1\right)$ gives

$$\lambda_t = 1 - q_t + \frac{k_{t+1} - k_t}{v}. \tag{23}$$

Another first-order condition (derived with respect to K_{t+1}) is

$$q_t = \frac{1}{1+r}\left(E_t\left[R_{t+1}\right] - \frac{1}{2}\frac{1}{v}\left(\frac{I_{t+1}}{K_{t+1}}\right)^2 + E_t\left[q_{t+1}\right] - \omega E_t\left[\lambda_{t+1}\right]\right), \tag{24}$$

where R_{t+1} is the capital rental rate. Equation (24) can then be simplified as

$$(1+r)q_t = E_t\left[\pi - \rho k_{t+1} + a_{t+1}\right] + E_t\left[q_{t+1}\right] - \omega E_t\left[\lambda_{t+1}\right]. \tag{25}$$

Substituting $E_t[\lambda_{t+1}]$ in equation (25) with equation (23) , we get

$$(1+r)q_t = E_t[\pi - \rho k_{t+1} + a_{t+1}] + E_t[q_{t+1}] - \omega E_t\left[\frac{k_{t+2}-k_{t+1}}{v} + 1 - q_{t+1}\right]. \tag{26}$$

In the constrained regime,

$$K_{t+2} = (1+\omega)K_{t+1} + W_t. \tag{27}$$

Without the loss of generality, we assume $W_t = 0$ at the moment, then

$$k_{t+2} = \ln(1+\omega) + k_{t+1},$$

where k_{t+1} is the log of K_{t+1}. Equation (26) can then be rewritten as

$$(1+r)q_t = \pi - \rho k_{t+1} + E_t[a_{t+1}] + E_t[q_{t+1}] - \omega E_t\left[\frac{\ln(1+\omega)}{v} + 1 - q_{t+1}\right]. \tag{28}$$

To solve for Tobin's q_t', we assume that

$$q_t' = B_0' + B_1' a_t + B_2' k_t. \tag{29}$$

By plugging equation (29) into equation (28), we have

$$\begin{aligned} & (1+r)\left(B_0' + B_1' a_t + B_2' k_t\right) \\ = \; & \pi - \rho(\ln(1+\omega) + k_t) + \gamma a_t + B_0' + B_1'\gamma a_t + B_2'(\ln(1+\omega) + k_t) \\ & -\omega\left(\frac{\ln(1+\omega)}{v} + 1 - \left(B_0' + B_1'\gamma a_t + B_2'(\ln(1+\omega) + k_t)\right)\right). \end{aligned} \tag{30}$$

By comparing coefficients, we get

$$B_0' = \frac{\left(\omega^2 - r\omega - v\rho - rv\rho\right)\ln(1+\omega) + v(r-\omega)(\pi-\omega)}{v(r-\omega)^2}$$
$$B_1' = \frac{\gamma}{1+r-\gamma-\gamma\omega}$$
$$B_2' = \frac{\rho}{\omega-r}.$$

Then, as ω rises, B_1' will increase, and hence the volatility of q_t' will also increase (given a shock to the technology a_t). That is, when the constraint always binds, inferior creditor protection will reduce stock price volatility. The intuition is that a credit constraint would reduce the upside potential of good shocks of A_{t+1} by constraining firm growth.

Note also that as ω rises, the level of Tobin's q_t' may go up or down. On the one hand, because $\partial B_1'/\partial\omega > 0$, higher ω allows firms to utilize the upshot of technology shock and raise firm value. But on the other hand, $\partial B_2'/\partial\omega < 0$. The intuition is that as ω rises, firms are able to invest more due to the increased investment ceiling ωK_t, which then increases K_{t+1} and reduces the capital rental rate R_{t+1} and hence Tobin's q_t'. The overall sign of $\partial q_t'/\partial\omega$ will then depend on the relative size of $\partial B_1'/\partial\omega$ and $\partial B_2'/\partial\omega$.[5]

2.3 Volatility of Tobin's q

So far we have looked at the variance of Tobin's q under the binding regime (equation (29)) and the non-binding regime (equation (15)) respectively. However, the investment may switch between binding and nonbinding regimes depending on the realization of an exogenous aggregate shock, W_t. Recall that there are two random variables: aggregate productivity shock ε_t and aggregate liquidity shock W_t. That is, we allow for aggregate uncertainty unrelated to liquidity shocks in addition to the liquidity shocks, and will flash out the distinction between productivity and liquidity aggregate shocks. Assume ε_t and W_t are independent. Then, by variance decomposition, we can write the

[5] Note that even if the permanent shock $W_t \neq 0$, the above qualitative results will still apply in the constrained regime, but it will be more complicated to write down a closed-form solution for q_t.

overall variance of Tobin's q as

$$Var\left[q_t\right] = E\left[Var\left[q_t|U_t\right]\right] + Var\left[E\left[q_t|U_t\right]\right], \tag{31}$$

where U_t is a dummy indicator for the binding regime. That is, $U_t = 1$ when the constraint binds.

Note that

$$E\left[Var\left[q_t|U_t\right]\right] = \Pr\left(U_t = 0\right) Var\left[q_{t,unconstrained}|U_t = 0\right] + \Pr\left(U_t = 1\right) Var\left[q_{t,constrained}|U_t = 1\right].$$

Combining equations (15) and (29), we have

$$E\left[Var\left[q_t|U_t\right]\right] = \left(\Pr\left(U_t = 0\right) B_1^2 + \Pr\left(U_t = 1\right) B_1'^2\right) Var\left[\varepsilon_t\right].$$

Moreover,

$$Var\left[E\left[q_t|U_t\right]\right] = \Pr\left(U_t = 1\right)\left(1 - \Pr\left(U_t = 1\right)\right)\left(\bar{q}_{t,unconstrained} - \bar{q}_{t,constrained}\right)^2,$$

where $\bar{q}_{t,constrained}$ stands for $E\left[q_t|U_t = 1\right]$ and $\bar{q}_{t,unconstrained}$ for $E\left[q_t|U_t = 0\right]$.

The effect of ω on $Var\left[q_t\right]$ is quite complex, as ω appears as an argument in the various expressions including $\Pr\left(U_t = 1\right)$, B_1' and $\bar{q}_{t,constrained}$. Take the extreme case where the technology is constant, i.e., $Var\left[\varepsilon_t\right] = 0$, we have

$$\begin{aligned} Var\left[q_t\right] &= Var\left[E\left[q_t|U_t\right]\right] \\ &= \Pr\left(U_t = 1\right)\left(1 - \Pr\left(U_t = 1\right)\right)\left(\bar{q}_{t,unconstrained} - \bar{q}_{t,constrained}\right)^2, \end{aligned} \tag{32}$$

and

$$\frac{\partial Var\left[q_t\right]}{\partial \omega} = \left(1-2\Pr\left(U_t=1\right)\right)\left(\bar{q}_{t,unconstrained}-\bar{q}_{t,constrained}\right)^2 \frac{\partial \Pr\left(U_t=1\right)}{\partial \omega} + \Pr\left(U_t=1\right)\left(1-\Pr\left(U_t=1\right)\right)\frac{\partial\left(\bar{q}_{t,unconstrained}-\bar{q}_{t,constrained}\right)^2}{\partial \omega}. \tag{33}$$

Now let us examine how each component of equation (33) is affected by the creditor protection parameter ω.

To see how the credit-regime switch probability, $\Pr\left(U_t=1\right)$, is related to ω, compare the optimal level of investment in the absence of credit constraint, $I_{t0} = vK_t\left(B_0+B_1a_t+B_2k_t-1\right)$, with the expression $\left(\omega K_t - W_t\right)$:

$$\Pr\left(U_t=1\right) = \Pr\left(I_{t0} > \omega K_t - W_t\right)$$

The expression $\Pr\left(I_{t0} > \omega K_t - W_t\right)$ negatively depends on ω, hence $\partial \Pr\left(U_t=1\right)/\partial\omega < 0$. That is, a reduction of ω is likely to increase the probability of a switch to the binding regime. Meanwhile, $\left(\bar{q}_{t,unconstrained}-\bar{q}_{t,constrained}\right)^2$ may depend positively or negatively on ω. This is because, as shown above, q_t' may rise or fall with ω, depending on the relative size of $\partial B_1'/\partial\omega$ and $\partial B_2'/\partial\omega$.

Our main proposition is that the overall effect of ω on $Var\left[q_t\right]$ is likely to be negative (through its effect on $\Pr\left(U_t=1\right)$), for a large class of distribution functions for W_t.[6] That is, an improvement in creditor protection lowers the stock market variability, through the reduction in the probability of credit-regime switch and market price oscillations across the two regimes.

Moreover $\left(\bar{q}_{t,unconstrained}-\bar{q}_{t,constrained}\right)^2$ may decrease as ω rises, particularly when $\partial B_1'/\partial\omega$ dominates $\partial B_2'/\partial\omega$, which will then reduce the impact of $\partial \Pr\left(U_t=1\right)/\partial\omega$ on $\partial Var\left[q_t\right]/\partial\omega$. Hence another prediction of the model is that by strengthening creditor protection, regulatory authorities may make the correlation between the regime-switch probability and the volatility of stock prices less pronounced. We will confront these predictions with cross-country panel data in the next

[6] The sign of $\left(1-2\Pr\left(U_t=1\right)\right)$ is likely to be positive, especially in the following empirical estimation where we associate the binding regime ($U_t = 1$) with the liquidity crisis, which has a sample probability of 10%.

section.

As a final note to the theoretical part, we want to mention that there is a simplification in our model. In the discussions on q_t' and $Var[q_t]$, we assume that the regime switch is permanent and occurs with a probability in period t. In a more general case where oscillations can take place in any future period, we can have the following Bellman equation

$$V(K_t, A_t, U_t; \omega) = \max_{Z_t, K_{t+1}} \left\{ A_t K_t^{1-\rho} - Z_t + \frac{1}{1+r} \sum_{U_{t+1}} \Pr(U_{t+1}|U_t) V(K_{t+1}, A_{t+1}, U_{t+1}; \omega) \right\},$$

where there is a Markov structure for the credit-constrained regime indicator U_t. The variance of firm's value, $Var[V(K_t, A_t, U_t; \omega)]$, will depend on the probability of the regime switch, $\Pr(U_{t+1}|U_t)$. With higher ω, we can expect higher $\Pr(U_{t+1} = 0|U_t = 0)$ and $\Pr(U_{t+1} = 0|U_t = 1)$, and lower $\Pr(U_{t+1} = 1|U_t = 0)$ and $\Pr(U_{t+1} = 1|U_t = 1)$. As higher ω reduces the probability of the switch from $U_t = 0$ to $U_{t+1} = 1$, we would still expect the overall effect of ω on $Var[V(K_t, A_t, U_t; \omega)]$ is negative for a large class of distribution functions for W_t.

3 Empirical Analysis

In this section we test whether credit constraints and creditor protection indeed affect stock market volatility in the way predicted by the model. We first discuss the data, empirical regularities, and the empirical approach and then present the results.

3.1 Data

The data used in this project come from the combination of sources as described in the Appendix.

Our creditor protection index comes from La Porta, et al. (1998).[7] The creditor rights index ranges from 0 to 4 with a higher number associated with better protection for creditors. The index is formed by adding one for each of the following four institutions: when the country imposes

[7] See http://post.economics.harvard.edu/faculty/shleifer/Data/l&fweb.xls.

Figure 1: The distribution of countries over creditor rights index (CR)

	Developing	Developed
CR=0	Colombia, Mexico, Peru, Philippines	France
CR=1	Argentina, Brazil	Australia, Canada, Finland, Greece, Ireland, Portugal, Switzerland
CR=2	Chile, Turkey	Belgium, Italy, Japan, Netherlands, Norway, Spain, Sweden
CR=3	Korea, South Africa, Thailand	Austria, Denmark, Germany, New Zealand
CR=4	China, Egypt, Hong Kong, India, Indonesia, Israel, Malaysia, Pakistan, Singapore	United Kingdom

restrictions, such as requiring a firm to obtain creditor consent or pay minimum dividends to file for reorganization; when secured creditors are able to gain possession of their security as soon as the reorganization petition has been approved (with no automatic stay); when secured creditors are ranked first in the distribution of the proceeds that result from the disposition of the assets of a bankrupt firm; and when the debtor does not retain the administration of its property pending the resolution of the reorganization. Figure 1 shows the countries in our sample that fall into different categories of the creditor rights index.

The data for stock market indexes come from Global Financial Data. We use monthly data calculated by central banks, national statistical agencies, or stock exchanges themselves as of the end–of–month closes. The country coverage includes emerging economies as well as developed economies for the years 1984-2004.[8] We convert all stock market indexes to U.S. dollars by multi-

[8]The panel is unbalanced.

plying them by the end–of–month exchange rates and scaling them down by the U.S. CPI at the end of the month.

To measure the stock return volatility (σ), we use the Officer (1973) method. This method estimates the stock return standard deviation for month 1 to month 12, then estimates the standard deviation from month 2 to month 13, and then repeats the procedure, continually rolling the sample forward. A potential problem with Officer's approach is that the use of overlapping observations can create a correlation between standard deviations at different points in time. An alternative is to use non–overlapping observations, that is, to compute the standard deviation using, say, months 1 through 12, 13 through 24, and so forth. Unfortunately, this procedure results in relatively few data points. We tried both methods and obtained similar results.

The rest of the variables used in the analysis are described in the Appendix.

3.2 Empirical Regularity

Figure 2 demonstrates the link between creditor protection and stock return variability. We group the values of the creditor rights protection index (CR) into high (3,4) and low (0,1,2) so that our results are not influenced by individual countries.[9] We can see from the Figure that better creditor protection is associated with lower stock price volatility. This relationship is confirmed statistically: the linear regression of the log of stock return volatility (σ) on the indicator of a high level of credit rights protection (CRH), developed country dummy (DEV), and the interaction of the two yields the following result:

$$Log(\sigma) = 2.33 - 0.59 * DEV - 0.42 * CRH + 0.29 * DEV * CRH + \varepsilon,$$

where ε is a robust standard error. All coefficients are statistically significant at 1–percent confidence level, the total effect of CRH for developed countries is significantly negative at the 3–percent

[9] We repeat all the results below with both the raw index and the full set of five indicators for each value of the index. As we report in the robustness tests section, our results are not affected qualitatively by these modifications.

confidence level, adjusted R^2 is equal to 0.21, and 788 observations are used. The magnitude of the effect of creditor rights on stock market volatility is non–negligible, although not very large — an increase in creditor protection from low to high for an average developing country would lower its stock market volatility by 80 percent of the standard deviation; for an average developed country, it would lower volatility by about a quarter of the standard deviation.

3.3 Empirical Approach

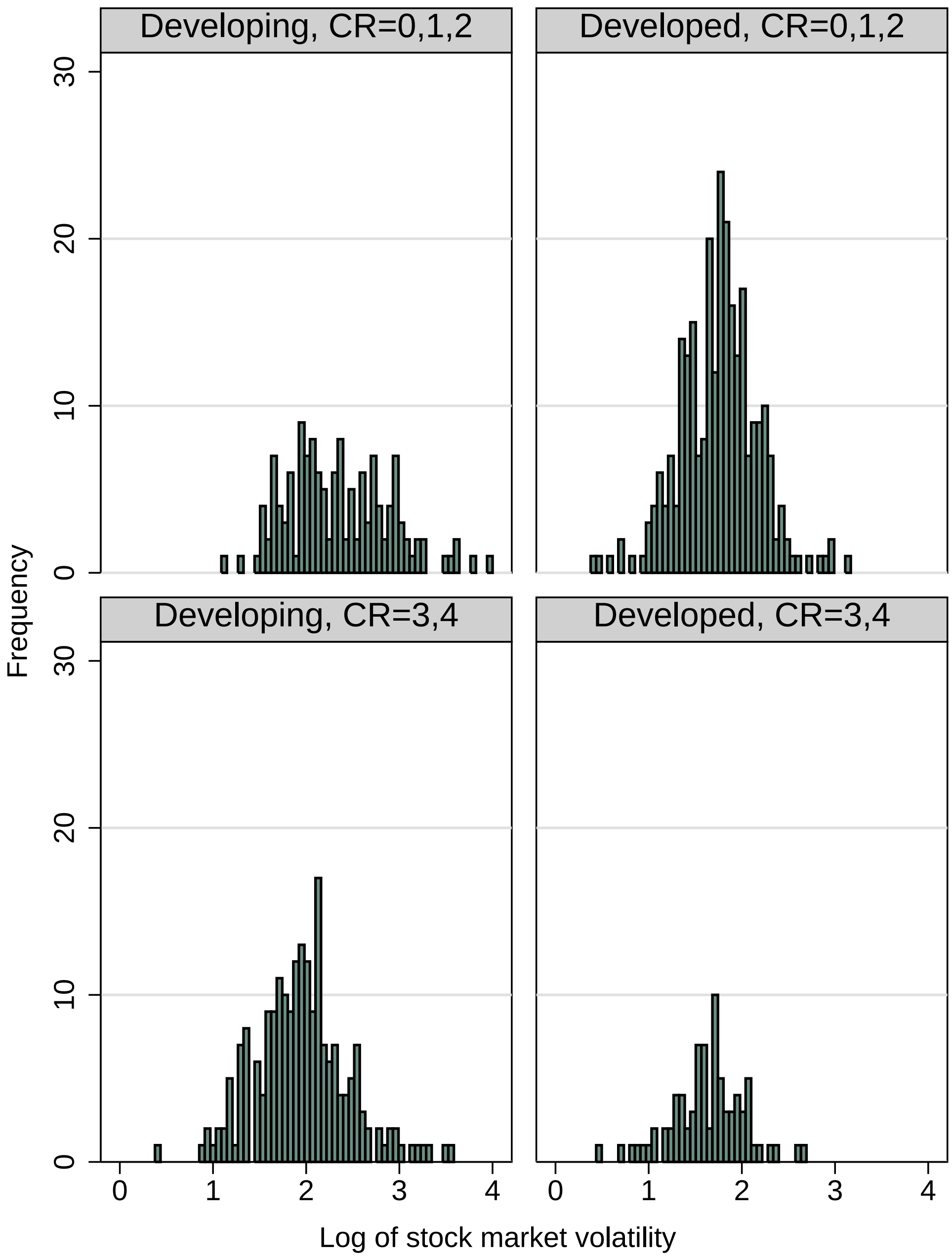
In our theoretical model, the credit constraint mechanism works through a random situation, where the constraint moves between binding and nonbinding. That is, the mechanism is based on a probability that the credit constraint is binding. In the empirical model, we use the probability of a liquidity crisis to proxy for the probability of binding. Hence, our empirical measure of the liquidity crisis is directly related to the theoretical counterpart of the credit constraint.

We define a liquidity crisis as an increase in the real interest rate of over 8.42 percentage points in one year, which corresponds to the highest 5 percent of annual changes in real interest rate in our sample.[10] We also define an alternative measure, to be used for the robustness tests, where a crisis is defined as an increase in the real interest rate of over 4.28 percentage points in one year, or the top 10 percent of annual real interest rate changes. Table 1 presents a list of countries and years for which our measure indicates a liquidity crisis has occurred and thus the indicator is equal to 1. Thus, our liquidity crisis variable measures domestic liquidity crises and proxies for the times when credit constraints are likely to be binding.

Following the methodology in Razin and Rubinstein (2006), we use our liquidity crisis indicator

[10]We obtain the data on interest rates from IMF International Financial Statistics. In most cases we use the money market rate. When the money market rate is not available, we use the discount rate. We calculate the real interest rate by subtracting the inflation rate from the nominal interest rate. We then calculate annual percentage changes in these real interest rates to identify liquidity crisis episodes as described in the previous section.

Figure 2: The distribution of stock market volatility.



to estimate the following model:

$$\mathrm{I(crisis)}_{it} = \begin{cases} 1 & \text{if} \quad y_{it} > 0 \\ 0 & \text{if} \quad y_{it} \leq 0 \end{cases},$$

where y is a latent variable and a function of our independent variables

$$y_{it} = X'_{it}\beta + \varepsilon_{it},$$

and ε has either a normal or a logistic PDF. X'_{it} includes the indicator of the political situation in the country (POL), as measured by the ICRG political risk index; a measure of capital mobility (CAP); and the lagged liquidity crisis indicator. We then construct a measure of the probability of a liquidity crisis (PLC) as a predicted value from the above estimation, which we use in the analysis of stock market volatility.

In the second stage, we measure the effects of credit constraints and of their interaction with the creditor rights measure. Specifically, we estimate

$$\ln(\sigma_{it}) = \alpha_i + \gamma_1 * PLC_{it} + \gamma_2 * (PLC_{it} * CRH_i) + Z'_{it}\delta + \eta_{it},$$

where $\ln(\sigma_{it})$ is our measure of the stock market volatility, for December of each year, $\{\alpha_i\}$ are country fixed effects, CRH_i is a time–invariant indicator of strong creditor rights protection, and errors η_{it} are allowed to be serially correlated and heteroskedastic. Z_{it} is a set of control variables, including the size of the stock market measured by the log of the number of listed firms, the growth rate of GDP per capita, and *de jure* financial account openness from Edwards (2006). Our theory guides us to distinguish between the effects of liquidity and productivity shocks on stock price volatility. That is why we include above two group of determinants of the volatility: a) those associated with productivity-based shocks, such as the GDP growth rate; and b) those associated with the probability of liquidity shocks, i.e, PLC_{it} and its interaction with CRH_i.

Evidently, one cannot possibly explain all the cross–country differences that would affect the variations in the stock market volatility between countries by institutional variables. Thus, we employ country–specific fixed–effects regression analysis, estimating the above equation by iterated feasible GLS (FGLS) with AR(1) disturbances. Note that since our creditor rights measure does not vary over time, it drops out from these regressions. Moreover, because the level of financial development varies vastly across countries, we believe the determinants of stock market volatility may vary as well. Thus, we estimate the second–stage regressions for the subsamples of developed and developing countries as well as for the full sample.

The above two–stage system can be identified with any set of explanatory variables through functional form. However, functional form identification tends to be weak, which is why we include in the first stage the variables that are likely to affect the probability of liquidity crisis but that do not have a direct effect on stock market volatility. In the first stage we use as an instrument the indicator of liquidity crisis lagged by one year. Because stock market prices tend to be forward–looking and efficient in processing information, past liquidity crises are unlikely to have a *direct* effect on the volatility of the stock market, although they are likely to affect the probability of future crises.[11] We test this exclusion restriction informally and find that, in the regressions of the stock market volatility, where we include a lagged dependent variable and there is no remaining autocorrelation in errors, the past liquidity crises indeed do not have any explanatory power (Table 2). Thus, we identify this system by both functional form and exclusion restriction.[12] In addition, we experiment with additional lags of liquidity crises and find that, while the fit of the first stage improves, the second–stage results are unaffected. Moreover, we reestimated our model using the Arellano–Bond (1991) dynamic panel technique and found the same results.

[11]See Fama (1991) for empirical evidences of the weak-form efficient market hypothesis.

[12]Note that, according to Table 2, the political risk index can also be interpreted as an instrument, since it does not appear to have a direct statistical effect on the stock return volatility. This helps us identify the model. However, we do not rely heavily on this index as an instrument because we do not have theoretical grounds to believe in its exogeneity.

3.4 Empirical Results

We now report the results of the two–stage estimation procedure: probability of crises and stock price volatility. Here we report the results of our analysis using the more strict definition of a liquidity crisis. We estimated all the models with a less–strict definition and found that our results are very similar, although the coefficients of interest in the second stage are smaller in magnitude.

3.4.1 Probability of Liquidity Crises

In the estimation of the first stage, we find that

$$\Pr(crisis) = 1.36 - 1.10 * CRH - 0.03 * POL - 0.01 * CAP + 0.76 * crisis_{-1},$$

Recall that POL is the indicator of the political situation in the country, and CAP is a measure of capital mobility. We find that crises are persistent and that better creditor protection, a more stable political situation, and more open financial accounts all lower the probability of a domestic liquidity crisis. All coefficients are significant at 1-percent confidence level. The McFadden's adjusted R^2 for this regression is 0.28 and 714 observations are used.[13] We use predicted values of this regression as a probability of liquidity crisis, our proxy for the tightness of the credit constraints, in the second stage.

The distribution of predicted crisis probability is shown for developed and developing countries by the level of credit protection in Figure 3.[14]

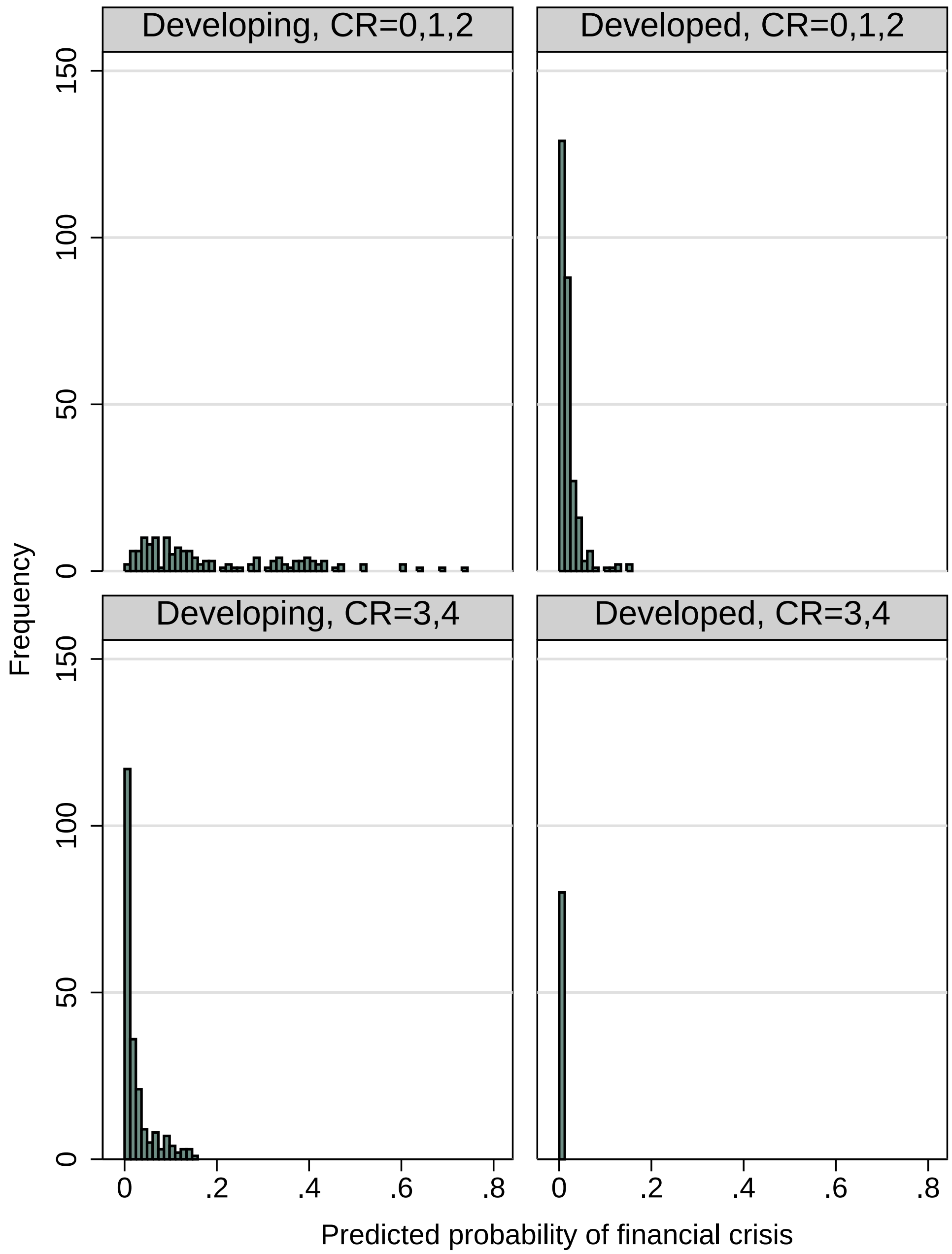
3.4.2 Stock Market Volatility

Table 3 reports the results of our second–stage estimation with country fixed effects. We include a lagged dependent variable and control for the growth rate of GDP per capita, for the log of

[13]The described regression is estimated using probit. We obtain almost identical results with logit regressions. Other conrol variables do not enter significantly and do not affect the results, thus we dropped them from the regression.

[14]Including a developed country dummy in the probit regression has only minor effect on this distribution.

Figure 3: The distribution of predicted probability of liquidity crisis



the number of firms listed on the stock market, and for the financial account openness. The table presents the results for the full sample (columns (1), (4), and (7)), the developing countries (columns (2), (5), and (8)), and the developed countries (columns (3), (6), and (9)).

The model features the probability of next period liquidity crisis as a predictor of the next period stock price volatility. Guided by this theory we perform a two–stage estimation procedure as described above. In the first stage we estimate the crisis probability by using the probit method. A key explanatory variable in the probit equation is the high creditor rights protection indicator. In the second stage we regress the stock price volatility on the predicted probability of liquidity crisis, the index of high creditor rights protection, and other covariates.

As the model predicts, our proxy for the tightness of the credit constraint, the predicted probability of liquidity crisis, increases stock market volatility. The effect is larger for developed countries, because the average predicted probability of liquidity crisis in developed countries is much lower (see Figure 3). To test the prediction of the model that weak creditor protection makes the effect of the crisis probability on the stock market volatility more pronounced, we include an interaction term between the indicator of the high credit rights protection index and the predicted crisis probability. We find that, as the model predicts, the effect of tight financial constraints on stock market volatility is mitigated for countries with good creditor rights protection — in fact, the coefficient on the interaction term more than compensates for the negative effects of the crisis probability.[15] According to the F-test, there is no significant effect of the crisis probability on stock market volatility for the countries with a high creditor rights protection index (the P-value of the F-test is 0.12 for the full sample).

Regarding the magnitudes of the effects, the average probability of liquidity crisis in the sample is 0.05 with a standard deviation of 0.10. Thus, the coefficient in column (4) of Table 3 implies that an increase in the probability of liquidity crisis by one standard deviation is associated with

[15]The coefficient on the interaction term in column (6) is very large and noisily estimated because the probability of liquidity crisis is very small for the developed countries with high level of creditor rights protection — see Figure 3.

a 6.6 percent increase in the stock market volatility in the countries with poor creditor rights protection.[16]

Thus, our empirical analysis confirms the effect of creditor protection on the correlation between the probability of liquidity crisis and the degree of stock market volatility in an economy with credit constraints: poor creditor protection increases the probability of tight financial conditions, and tight financial conditions are found to increase stock market volatility through the crisis–probability channel. Moreover, better creditor protection mitigates the effects of tight financial conditions on stock market volatility, consistent with the predictions of the model. While the effects we find are not very large, they are likely to be biased downwards because of the measurement error associated with our proxy for credit constraints, as well as the measurement error in the *de jure* definition of the creditor rights index.[17]

Note that our empirical approach focuses on the case where the constraint shifts between binding and nonbinding regimes. Ideally, we would also hope to have a variable to proxy for the scenario when the constraint always binds, but it is hard to find such a macroeconomic variable. Hence, we use country fixed effects α_i to control for the case when the constraint always binds. Doing so can give us unbiased estimation of the effect of creditor rights on price volatility through affecting the probability of regime switching.

3.4.3 Robustness Tests

We conduct a series of robustness tests to make sure our findings are not driven by the exact specification we have chosen. We describe them in this section but do not report the regression tables in the interest of space. The tables are available from the authors upon request.

We attempted using additional control variables, such as the fiscal situation in the country,

[16] The log of stock market volatility increases by 0.0642, which means stock market volatility itself increases by 6.6 percent. While this number is economically meaningful, it is not very large, given that the standard deviation of our stock market volatility measure is about 70 percent of its mean for developing and 45 percent for developed countries.

[17] It is well known that while the creditor rights index takes on a value of 4 in countries like India and China, *de facto* creditor protection in these countries is low.

current account, stock market price/earnings ratio, fixed and floating exchange rate regime indicators (Reinhart & Rogoff, 2004), and volatility of the U.S. 3-year Treasury–bill rate, but none of these variables entered the regressions with significant coefficients or affected the results in any way, save for some of them limiting the sample. Sovereign credit rating does enter significantly in the regressions, but it is highly correlated with the growth rate of GDP per capita (with the correlation coefficient of 0.79), which is why we did not include it in the main specification.

We reestimate our first–stage regression using two or three lags of liquidity crisis. The R^2 of the probit regression increases and all the lags have positive and significant effects. The results of our second stage are not qualitatively affected. Quantitatively, the coefficients on the predicted crisis probability increase in magnitude, while the coefficients on the interaction term lose precision as we lose observations due to the introduction of the lags: while the interaction effects are still statistically significant with two lags of liquidity crisis indicator in the first stage, with three lags the P-value rises to 0.11-0.19.

We repeat our analysis with an alternative (less–strict) definition of the liquidity crisis in the first stage, which leads to a different predicted probability of crisis. The correlation between our old and new predicted crisis probability is very high: 0.89. We repeat our second–stage estimation with this new crisis probability and find no qualitative differences in our results. Quantitatively, the effects of crisis probability are smaller, as expected, but are still statistically significant.

Going back to our original definition of liquidity crisis, we now use a logit model to construct our predicted crisis probability. The correlation of the new measure with the original one is again very high: 0.99. We reestimate Table 3 using this new prediction. As expected, given the high correlation of the measures of crisis probability, the estimated coefficients are almost identical to our main specification. Alternatively, we included a developed country dummy in our probit regression, which left our results unchanged.

Instead of using a binary indicator for a high level of creditor rights protection, we used a raw index as if it were a continuous variable and, alternatively, a set of five dummy variables, one for each value of the index. Our results are qualitatively the same, with most of the effect of high level of creditor rights protection appearing in the coefficient on the dummy variable indicating that the creditor rights protection index is equal to 4.

We reestimated the regressions in Table 3 excluding the lagged dependent variable. We find that our results still hold and that the coefficient estimates are now slightly larger and are still significant. As we would expect, when the lagged dependent variable is excluded, there is serial correlation in the error term: the AR(1) coefficient is about 0.2. We also reestimated the regressions in Table 3 using the Arellano-Bond dynamic panel model and found that the results are qualitatively the same.

In the estimation of Table 3, we did not correct our standard errors for using the predicted probability as an explanatory variable. As Heckman (1978) points out, consistent estimates of variance can be obtained if the predicted probability is used as an instrument for the binary variable on the right-hand side. We reestimated our model in this way (with and without fixed effects), using GMM, and found that our results are robust to this correction.

We repeated our analysis adding year fixed effects. We found that our effects are qualitatively the same — all the coefficients that were initially significant are still significant and have the same sign. The only important quantitative effect now is a smaller coefficient on the crisis probability in all regressions (about half the size of those reported in Table 3).

Finally, instead of classifying countries into developed and developing, we reclassified them into OECD and non–OECD, which affected Mexico, Turkey, and Korea. In the case of Mexico and Korea, we only reclassified them for the years they were actually in the OECD. Our results remained qualitatively the same, indicating that none of these three countries strongly influences our results.

4 A Concluding Remark

In this paper, we examine the connection between creditor rights protection and the volatility of stock market prices. Obtaining an analytical solution in a version of a Tobin's q investment model with credit constraints allows us to focus on the predictions regarding the second moments of q. We find that better creditor protection, and hence lower collateral requirements, may reduce the price volatility through two channels: by lowering the probability of switching to the constrained state and by lowering the stock price volatility associated with this switch.

Using a panel of 20 developed and 20 developing countries for the last two decades, we find empirical support for the predictions of our model: weak creditor protection increases the probability of liquidity crises, which is our proxy for the probability of the regime with binding credit constraint, and hence increases the aggregate stock price volatility. Moreover, the second–stage effect of the probability that the constraint is binding on the stock return volatility is mitigated in the countries with good creditor protection.

Our paper thus illustrates the importance of creditor protection on the development of a sound stock market: strong creditor rights not only increase the stock value, but also crucially, reduce the counterproductive volatility of the stock market. This finding is relevant for the recent credit crunch in developed markets that was associated with high stock market volatility. While Germany was the country most affected by the liquidity crisis, the stock market volatility increase was less pronounced in Germany than it was in France, Australia, or Japan, which all have a lower degree of creditor rights protection.

Finally, there are other mechanisms through which creditor protection may affect the volatility of stock market prices. For instance, Hale, Razin, and Tong (2006) discuss the moral hazard channel. Weak creditor protection induces firms to make riskier investments, as firms will benefit from the upper range of the realized capital return, with no need to worry about the lower range. Such moral hazard can increase stock price volatility. We leave it to future work to test this prediction.

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Appendix

In the regressions that are reported we used the data series constructed from the variables listed below. In our robustness tests we used a host of additional control variables that were obtained mostly from the IFS and the Global Financial Data.

Variable	Units	Frequency	Source
Creditor rights index	Index 0-4	cross-section	La Porta, et al. (1998)
Composite stock market close	Index	monthly (eop)	Global Financial Data
Exchange rate against U.S. dollar	n.c./U.S.dollar	monthly (eop)	Global Financial Data
U.S. CPI	Index	monthly (eop)	Global Financial Data
Deposit rate	percent	annual	IFS, line 60l
Money market rate	percent	annual	IFS, line 60b
Inflation rate	percent	annual	IFS, line 64..x
GDP in U.S. dollars	millions of USD	annual	Global Financial Data
Population	thousands of people	annual	Global Financial Data
De jure financial account openness	Index 0-100	annual	Edwards (2006)
Exchange rate regime	Index 1-6	annual	Reinhart & Rogoff (2004)
Index of political stability	Index 0-100	annual	ICRG
Companies listed on stock markets	units	annual	Global Financial Data
U.S. 3-year T-bill rate	percent	monthly	FAME

Table 1: List of liquidity crises in the sample

Country	Years of financial crisis
Argentina	1984, 1987, 1988, 1989, 1990, 1992, 1993[a], 1994[a], 2001, 2004[a]
Australia	1984[a], 1989[a]
Brazil	1987, 1988, 1989, 1990, 1992, 1993, 1994, 1996, 1997[a], 1998[a]
Chile	1984[a], 1987[a], 1989
China	1990[a], 1995[a], 1996[a]
Colombia	1998
Egypt	1985[a], 1990[a], 1992[a], 1996[a]
Greece	1987[a], 1988[a]
Hong Kong	1999[a]
India	1984[a], 1989[a], 1995[a]
Indonesia	1984[a], 1997
Israel	1984, 1986, 1987, 1992[a], 2003[a]
Korea	1989[a]
Mexico	1984, 1985, 1989, 1995, 1998
Peru	1991, 1992, 1993, 1995[a], 1999[a]
Philippines	1985, 1986, 1992, 1997[a]
Portugal	1985[a], 1991[a]
South Africa	1984[a], 1988[a]
Spain	1987[a]
Sweden	1992
Thailand	1997[a]
Turkey	1990, 1991, 1994, 1996, 1998[a], 1999, 2001, 2003[a]

[a] No liquidity crisis by our strict definition

Table 2: Informal tests of exclusion restrictions

	Full Sample (1)	Developing (2)	Developed (3)
Lagged dependent variable	0.35***	0.37***	0.29***
	(0.036)	(0.046)	(0.054)
Growth rate of GDP per capita	-0.57***	-0.88***	-0.062
	(0.11)	(0.14)	(0.19)
Log (# of firms listed on the stock market)	0.009	-0.053	-0.083
	(0.038)	(0.052)	(0.074)
Financial account openness	-0.003***	-0.003*	-0.003
	(0.001)	(0.002)	(0.002)
ICRG political risk index	-0.000	-0.002	0.007
	(0.002)	(0.003)	(0.005)
Lagged liquidity crisis indicator	0.096	0.089	0.251
	(0.064)	(0.064)	(0.319)
Observations	679	344	335
Number of group(country)	40	20	20
Log likelihood	-274	-145	-120
Common AR(1)	-0.035	-0.036	-0.013

Iterated FGSL. Standard errors in parentheses
Dependent variable is log of stock return volatility.
Country fixed effects are included
* significant at 10%; ** significant at 5%; ***significant at 1%

Table 3: Second–stage regressions

	Full Sample (1)	Developing (2)	Developed (3)	Full Sample (4)	Developing (5)	Developed (6)	Full Sample (7)	Developing (8)	Developed (9)
Lagged dependent	0.35***	0.38***	0.27***	0.34***	0.36***	0.26***	0.33***	0.36***	0.27***
variable	(0.035)	(0.046)	(0.054)	(0.035)	(0.046)	(0.054)	(0.035)	(0.046)	(0.054)
Growth rate of	-0.58***	-0.88***	-0.06	-0.56***	-0.86***	-0.010	-0.56***	-0.87***	-0.019
GDP per capita	(0.11)	(0.14)	(0.19)	(0.11)	(0.13)	(0.19)	(0.11)	(0.13)	(0.19)
Log (# firms listed	0.014	-0.067	0.12*	0.024	-0.046	0.16**	0.008	-0.075	0.16**
on the stock mkt.	(0.040)	(0.047)	(0.071)	(0.040)	(0.047)	(0.073)	(0.041)	(0.049)	(0.073)
Financial account	-0.004***	-0.004**	-0.003	-0.002	-0.001	-0.000	-0.002	-0.001	-0.001
openness	(0.001)	(0.002)	(0.002)	(0.001)	(0.002)	(0.002)	(0.001)	(0.002)	(0.002)
Predicted probability				0.59***	0.59***	3.89*	0.64***	0.62***	3.51*
of liquidity crisis (PLC)				(0.21)	(0.22)	(1.95)	(0.21)	(0.22)	(1.98)
CRH*PLC							-2.23**	-2.64**	-366.0
							(1.03)	(1.08)	(286.9)
Observations	679	344	335	679	344	335	679	344	335
Number of countries	40	20	20	40	20	20	40	20	20
Log likelihood	-275	-145	-121	-271	-143	-119	-270	-140	-119
Common AR(1)	-0.034	-0.034	-0.008	-0.033	-0.036	-0.006	-0.032	-0.036	-0.002

Iterated FGLS. Standard errors in parentheses
Dependent variable is log of stock return volatility.
Country fixed effects are included
* significant at 10%; ** significant at 5%; *** significant at 1%