THE RISKY SPREAD, INVESTMENT, AND MONETARY POLICY TRANSMISSION: EVIDENCE ON THE ROLE OF ASYMMETRIC INFORMATION

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Abstract—Financing constraints can arise when there are important information asymmetries in financial markets. Using Canadian panel data, we reject a symmetric information specification of investment behaviour in favour of an agency cost specification in which the shadow cost of finance can diverge from the market interest rate. Our empirical estimates suggest that shocks to net worth, as reflected in the risky spread and firm-specific balance sheet variables, can dramatically increase the shadow cost of finance. Tests which draw on distinctive institutional features of the Canadian economy show that it is firms in a weak informational position which tend to be responsible for this result.

I. Introduction

THE question of whether firms face finance constraints has important implications for macroeconomics. The issues have been discussed in recent work by Fazzari, Hubbard and Petersen (1988) among others. Finance constraints induced by capital market and information imperfections can raise the external cost of borrowing above the risk-free market interest rate. As discussed in Bernanke and Gertler (1990), Bernanke and Blinder (1988), and Kashyap, Stein and Wilcox (1993), this may provide an additional transmission mechanism for monetary policy which is not present under perfect capital markets. Monetary policy may affect output not just through the standard channel of the interest rate, but also through the wedge between the market and the shadow cost of finance. Furthermore, if the likelihood that finance constraints bind is cyclical, monetary policy may have asymmetric effects on output over the business cycle.

A number of recent papers have used Q investment equations to test for the importance of finance constraints on firms' investment spending,[1] but the approach has some drawbacks.[2] One problem is that Q equations relate investment to the expected stream of future marginal products of capital which are unobserved to econometricians. Liquidity variables like cash flow may be significant in Q equations simply because they capture information about investment opportunities that is omitted from the econometrician's information set. An alternative is to estimate Euler equations. Euler equations emphasize the period-to-period "arbitrage" decision as to whether to invest today or tomorrow and do not require the econometrician to form precise expectations of the distant future. The Euler equation approach has been extensively used in empirical studies of consumption and has recently been used by Hubbard and Kashyap (1992) and Whited (1992) in empirical investment work.

This paper uses Euler equations to analyze the importance of finance constraints from three angles. The first is based on the idea that the error term in the Euler equation reflects expectational errors and should be orthogonal to information available in period t under the null hypothesis of symmetric information and rational expectations. If finance constraints bind, the Euler equation should depend on the shadow cost of external financing. The symmetric information model which omits this information would be misspecified and the error term from estimating the symmetric information model will not be orthogonal to variables in the information set in period t.

While a rejection of the orthogonality conditions implied by the symmetric information model would be consistent with the finance constraints view, other forms of misspecification might also lead to a rejection. A more direct test for finance constraints would be to see if the shadow cost of finance is equal to the market interest rate. However, shadow costs are typically unobserved. Nevertheless, Gertler, Hubbard and Kashyap (1991) and Whited (1992) suggest estimating the Euler equation which takes into account finance constraints by parameterizing the shadow cost of finance. Our second test for the importance of finance constraints follows this route. We extend the work of Whited (1992) and Gertler et al. (1991) by allowing both aggregate and firm-specific balance sheet variables to affect the shadow cost of finance. Bernanke (1983) and Bernanke and Gertler (1989) suggest that the risky spread (defined as the difference between a risky interest rate and a riskless interest rate on securities of the same maturity) may reflect agency costs of financial intermediation. Gertler et al. (1991) build a simple optimal contracting model which shows why the wedge between the costs of internal and external finance might be linked to the risky spread. We therefore use the risky spread as an aggregate measure of agency costs. As surveyed in Gertler (1988), a number of authors have also suggested that balance sheet variables might be linked to agency costs. We follow Whited (1992) and focus on the debt-equity and interest coverage ratios as firm-specific measures of agency costs.

Our third set of tests takes explicit account that firms are heterogeneous. Firms which are more capable of credibly communicating private information to the capital market should be less affected by problems of asymmetric information. This implies that if agency costs are important, we should find stronger evidence against the symmetric information model for firms in a weak informational position. To test this hypothesis, we estimate Euler equations on samples separated according to variables which reveal firms' informational positions.

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[1] This literature includes Chirinko and Schaller (1995), Devereux and Schiantarelli (1990), Fazzari et al. (1988), Hoshi, Kashyap and Scharfstein (1991), Schaller (1993), and Whited (1991).

[2] Abel and Blanchard (1986) list some of the problems in aggregate time series estimates.

Our empirical analysis is based on a panel of Canadian data. The use of panel data reduces the severity of aggregation issues that typically arise with time series aggregate data,[3] and distinctive institutional features of the Canadian economy allow comparisons between firms which are differentially positioned to credibly communicate private information to outsiders. Most of the Euler equation tests for credit constraints on investment have been conducted on U.S. data. The Canadian evidence provides a check on the robustness of the U.S. results in a country that is broadly similar but in which the details of capital market structures differ.

The paper is organized as follows. Section II presents the derivation of the Euler equation for the symmetric and asymmetric information models. In section III we estimate and test the models on a panel of 199 publicly-traded Canadian firms over the period 1973–1986. Section IV concludes.

II. Derivation of the Empirical Specifications

We first consider the problem for a representative firm when there are no borrowing constraints. Firm-specific subscripts will be suppressed to avoid unnecessary notation. The objective of the firm is to maximize its value, V_0, as of period 0:

$$V_0 = E_0 \sum_{t=1}^{\infty} \left(\prod_{j=0}^{t-1} \beta_j \right) d_t,$$

where E_0 is the expectations operator conditional on information available at time 0, β_t is the discount factor at time t, or the inverse of one plus the appropriate discount rate, and d_t is dividends. The firm faces a capital accumulation constraint that

$$K_t = (1 - \delta)K_{t-1} + I_t,$$

where K_t is the capital stock at the end of period t, δ is the depreciation rate, I_t is investment. The firm also faces a non-negativity constraint on dividends,

$$d_t \geq 0 \ \forall t,$$

with d_t defined as

$$(1 - \tau_t)(\Pi(K_t, L_t) - G(I_t, K_t) - w_t L_t) - p_t^I I_t + B_t - (1 + r_{t-1})B_{t-1},$$

where $\Pi(K_t, L_t)$ is the revenue function, L_t is variable inputs, τ_t is the corporate tax rate, and w_t and p_t^I are the real price of variable inputs and investment, respectively. It is assumed that capital is costly to adjust, and $G(I_t, K_t)$ is a linear homogeneous function in I and K. The firm pays r_{t-1}, the after-tax real interest rate, on the stock of one-period debt outstanding at the end of period $t - 1$ and issues an amount B_t of new debt each period, subject to the transversality condition of a no Ponzi game that

$$\lim_{T \to \infty} \left(\prod_{t=0}^{T-1} \beta_t \right) B_T = 0.$$

Let λ_t^K and λ_t^d be the Lagrange multipliers on capital accumulation and the non-negativity constraint on dividends, respectively. Also, let H_x denote the partial derivative of the function H with respect to x. The first order conditions for capital, investment, and debt are, respectively:

$$(1 + \lambda_t^d)(1 - \tau_t)(\Pi_K(K_t, L_t) - G_K(I_t, K_t)) - \lambda_t^K + E_t \beta_t (1 - \delta)\lambda_{t+1}^K = 0, \tag{1}$$

$$\lambda_t^K = (1 + \lambda_t^d)(p_t^I + (1 - \tau_t)G_I(I_t, K_t)), \tag{2}$$

$$(1 + \lambda_t^d) - E_t[(1 + \lambda_{t+1}^d)\beta_t(1 + r_t)] = 0. \tag{3}$$

For future reference, we define

$$\tilde{\beta}_t = 1/(1 + r_t).$$

The first order conditions (1) to (3) imply

$$E_t((1 - \tau_t)(\Pi_K(K_t, L_t) - G_K(I_t, K_t) - G_I(I_t, K_t)) - p_t^I + (1 - \delta)\tilde{\beta}_t((1 - \tau_{t+1}) \times G_I(I_{t+1}, K_{t+1}) + p_{t+1}^I)) = 0.$$

In order to test the model, we must make assumptions about the functional forms. The revenue function is $\Pi(K_t, L_t) = P(Y_t)F(K_t, L_t)$, where $P(Y)$ is the inverse demand function, and $F(K_t, L_t)$ is the production function. We assume that the latter is homogeneous of degree $\eta = \alpha_L + \alpha_K$, viz:

$$F(K, L) = K^{\alpha_K} L^{\alpha_L} \tag{4}$$

with $F_i > 0$ and $F_{ii} < 0$, $i = K, L$, and $F_K K + F_L L = \eta F$. If the firm is a price taker, $\Pi_i = PF_i = F_i$, $i = K, L$, with the price of output as the numeraire. But if the firm faces a downward sloping demand curve, then $\Pi_i = (1 + 1/\epsilon)F_i$, where ϵ is the price elasticity of demand. Allowing for both non-constant returns to scale and imperfect competition, we have

$$\Pi_K = (1 + 1/\epsilon)\eta \frac{F}{K} - \frac{wL}{K} = \psi \frac{F}{K} - \frac{C}{K}$$

where C is the total cost of the variable factors of production, and $\psi = (1 + 1/\epsilon)\eta$. In the estimation, Π_K is replaced by

[3] Schaller (1990) shows that some, although not all, of the problems of empirical investment equations are due to aggregation.

the above expression. Note, however, that ϵ is not separately identifiable from η. Effects due to non-constant returns to scale and imperfect competition are all subsumed in the parameter ψ.

As is standard in the investment literature, the adjustment cost function is given by

$$G(I, K) = \frac{\phi}{2}\left(\frac{I_t}{K_t}\right)^2 K_t, \tag{5}$$

where ϕ is the marginal adjustment cost parameter. This implies $G_I = \phi I_t/K_t$, and $G_K = -(\phi/2)(I_t/K_t)^2$. Substituting in G_K and G_I, and assuming a rational expectations error of e_t satisfying $E_t(e_{t+1}) = 0$, we have

$$(1-\tau_t)\left(\psi\frac{Y_t}{K_t}-\frac{C_t}{K_t}+\frac{\phi}{2}\left(\frac{I_t}{K_t}\right)^2-\phi\left(\frac{I_t}{K_t}\right)\right)-p^I_t$$
$$+(1-\delta)\tilde{\beta}_t\left((1-\tau_{t+1})\phi\frac{I_{t+1}}{K_{t+1}}+p^I_{t+1}\right)=e_{t+1}. \tag{6}$$

Under the null hypothesis of rational expectations and that firms do not face borrowing constraints (i.e., under symmetric information), the error term e_{t+1} should be orthogonal to variables dated t. The Modigliani-Miller theorem applies in this case, and investment spending should be independent of the structure of financial claims on firms. This will not be the case in an asymmetric information model.

Under imperfect information, insiders within the firm can either invest or divert resources and appropriate the proceeds, but the lender can never be sure whether low output is the result of a negative shock or mismanagement. The optimal financial contract is structured to minimize the gap between repayment when the project fails and and when it succeeds, and thus the extent to which investment falls below its first-best level. The features of asymmetric information can be modeled as a debt capacity constraint on the firm. If B^*_t is the maximum amount of debt that the firm is allowed to issue, then

$$B_t \leq B^*_t. \tag{7}$$

Let ω_t be the Lagrange multiplier on the borrowing constraint. In place of (3), the first order condition for B_t now becomes:

$$(1+\lambda^d_t)-E[\beta_t(1+\lambda^d_{t+1})(1+r_t)]-\omega_t=0. \tag{8}$$

It is easy to see that when the debt capacity constraint does not bind, $\omega_t = 0$, then (8) reduces to (3), the case of symmetric information. Define $\tilde{\omega}_t = \omega_t/(1 + \lambda^d_t)$ and $\beta_t = E_t[((1 + \lambda^d_t)(1 - \tilde{\omega}_t))\ ((1 + \lambda^d_{t+1})(1 + r_t))^{-1}]$. It follows that when the constraint binds, the discount factor is smaller under asymmetric information than under symmetric information, ceteris paribus.

One way to think about the impact of financing constraints on investment behavior is that when a firm finds it difficult to obtain external financing, the market interest rate is a poor proxy for the shadow cost of external financing. The relevant cost to the firm may be much higher than the market interest rate, and the firm uses this higher rate to discount its net cash flows. A firm with lower net worth has less collateral and therefore a lower borrowing capacity. As suggested by Gertler et al. (1991), negative shocks to net worth may arise as a result of a decrease in the collateral firms can offer, or a disruption in credit markets which contaminates the information about firms gathered by financial intermediaries. Negative shocks to net worth can influence investment spending of finance-constrained firms by tightening the borrowing constraint. As the constraint is tightened, the shadow cost of finance rises and drives a wedge between the market interest rate and this shadow cost. As a result, the effective discount rate will be higher and may vary more than the market interest rate.

The Euler equation under asymmetric information is

$$(1-\tau_t)\left(\psi\frac{Y_t}{K_t}-\frac{C_t}{K_t}+\frac{\phi}{2}\left(\frac{I_t}{K_t}\right)^2-\phi\frac{I_t}{K_t}\right)$$
$$-p^I_t+\tilde{\beta}_t(1-\tilde{\omega}_t)(1-\delta)$$
$$\times\left((1-\tau_{t+1})\phi\frac{I_{t+1}}{K_{t+1}}+p^I_{t+1}\right)=e_{t+1},$$

which it is instructive to rewrite as:

$$(1-\tau_t)\left(\psi\frac{Y_t}{K_t}-\frac{C_t}{K_t}+\frac{\phi}{2}\left(\frac{I_t}{K_t}\right)^2-\phi\frac{I_t}{K_t}\right)-p^I_t$$
$$+\tilde{\beta}_t(1-\delta)\left((1-\tau_{t+1})\phi\left(\frac{I_{t+1}}{K_{t+1}}\right)+p^I_{t+1}\right)$$
$$=e_{t+1}+\tilde{\beta}_t\tilde{\omega}_t(1-\delta)\left((1-\tau_{t+1})\phi\frac{I_{t+1}}{K_{t+1}}+p^I_{t+1}\right). \tag{9}$$

The econometric implication of financing constraints is that if the asymmetric information model is the correct model, the Euler equation given by (6) will be misspecified. When $\tilde{\omega}_t \neq 0$, the borrowing constraint will induce an additional term in the Euler equation. Variables dated t will be correlated with the error term of the symmetric information model, and the data will reject the orthogonality restrictions. However, the Euler equation which takes the debt capacity constraint into account should hold if the asymmetric information model is correct.

To test these competing models, we need to parameterize $\tilde{\omega}_t$ since it is unobserved. We consider two specifications. The first is based on the risky spread, defined as the difference between a risky interest rate and a riskless interest rate on securities of the same maturity. The idea that the risky spread might capture agency costs of financial intermediation had earlier been suggested by Bernanke (1983) and Bernanke and Gertler (1989). Gertler et al. (1991) used it to parameterize the shadow cost of finance constraints in the Euler equation for the asymmetric information model and found a statistically significant relationship between the risky spread and aggregate investment in the United States. In our panel data context, the risky spread can be thought of as capturing the effect of aggregate shocks to internal net worth which might affect the shadow cost of finance. This suggests

$$\tilde{\omega}_t = \gamma_0 + \gamma_1 S_{t-1}, \tag{10}$$

where S_{t-1} is the risky spread.

The second specification is motivated by the link between agency costs and balance sheet variables. Using panel data, Whited (1992) found balance sheet variables to be statistically significant when used to parameterize the shadow cost of finance constraints. Recent work has found the risky spread to have considerable predictive power for aggregate output even in the presence of other economic leading indicators.[4] It is therefore of interest to see if the risky spread still has additional explanatory power for investment when firm-specific balance sheet information is also present. This suggests

$$\tilde{\omega}_t = \gamma_0 + \gamma_1 S_{t-1} + \gamma_2 X_{t-1}, \tag{11}$$

where X_{t-1} is a balance sheet variable such as the debt–equity ratio. The possibility that both aggregate and firm-specific factors might affect the severity of finance constraints has not been considered in previous studies. The above specification which adds $\gamma_2 X_{t-1}$ to the parameterization of $\tilde{\omega}_t$ is unique in relation to existing work in the literature.

Substituting (11) into (9) and rearranging terms gives the general specification for the agency cost model:

$$\begin{aligned}(1-\tau_t)&\left(\psi\frac{Y_t}{K_t} - \frac{C_t}{K_t} + \frac{\phi}{2}\left(\frac{I_t}{K_t}\right)^2 - \phi\frac{I_t}{K_t}\right)\\ &- p^I_t + \tilde{\beta}_t(1-\delta)(1-\gamma_0-\gamma_1 S_{t-1}-\gamma_2 X_{t-1})\\ &\times\left((1-\tau_{t+1})\,\phi\left(\frac{I_{t+1}}{K_{t+1}}\right) + p^I_{t+1}\right) = e_{t+1}.\end{aligned} \tag{12}$$

Equation (12) says that, after controlling for borrowing constraints, the error term in the Euler equation of the asymmetric information model should be orthogonal to variables dated t. A rejection of the orthogonality conditions for (6) in favour of those for (12) would be seen as support for the asymmetric information model.

[4] On the predictive power of the risky spread, see Stock and Watson (1989), Friedman and Kuttner (1992, 1993).

III. Results

Our analysis is based on a panel of 199 Canadian firms over the sample 1973 to 1986. At the macro level, the cyclical properties of the risky spread in Canada are similar to those of the United States, with the spread being largest during the recessions in the mid-1970s and early 1980s in both countries.[5] At the micro level, there are substantial variations in the characteristics of firms in Canada, just as in the United States. In spite of these similarities, there are institutional and structural differences which may cause firms in the two countries to behave differently. These will be analyzed in more detail below.

Our empirical analysis is based on the Generalized Method of Moments (GMM) approach of Hansen and Singleton (1982). Let N be the number of firms and T be the span of the data. The parameters are obtained by minimizing the objective function $g'\Omega g$, where $g = 1/(NT)\ \sum_{i=1}^{N}\sum_{t=1}^{T} h_{it}$, and h_{it} is the moment condition for firm i at time t. The weighting matrix, Ω, is the heteroskedastic-consistent long-run variance matrix constructed according to the method discussed in Newey and West (1987a).

Under the null hypothesis of rational expectations, the error terms in (6) and (12) are pure expectational errors and should be orthogonal to variables in the information set at t. These moment conditions are tested via the J-statistic for the overidentifying restrictions between a suitable set of instruments and the regression residuals. The data are first differenced prior to the estimation to remove fixed effects. First differencing induces a moving average component in the error term, but this serial correlation is accounted for in the construction of the weighting matrix. In view of this serial correlation, the instruments we use are values of the variables that appear in the Euler equation lagged two periods. For the base case, we use the $t-2$ values of Y/K, C/K, p^I, τ, δ, I/K, $(I/K)^2$, and S as instruments. Note that S is included as an instrument when estimating the symmetric information model to ensure that the instrument set is the same under both the null and alternative specifications. However, the weighting matrix for the symmetric information specification and the agency-cost specification is not identical. Unless otherwise stated, the t and J statistics are based on the weighting matrix of the model being estimated. We also constrain the adjustment cost coefficient ϕ to be positive for meaningful interpretation of the Euler equations.

A. *Testing the Symmetric vs. the Asymmetric Information Model: Full Sample Estimates*

The first row of table 1 presents estimates of the Euler equation under symmetric information. The estimates of ψ,

[5] The risky spread does not seem to have been as useful a leading indicator for the 1990–91 recession in the United States as it has been for previous recessions. In Canada, there was an increase in the spread before the 1990–1991 recession, but it was modest compared to the increases in the mid-1970s and early 1980s.

TABLE 1.—SYMMETRIC INFORMATION SPECIFICATION

	ϕ	ψ	J
	(a) Full Panel		
Risk-free rate	.682	.884	39.2
	(.212)	(.019)	[.0000]
Risky rate	.712	.881	43.3
	(.215)	(.019)	[.0000]
Alternative weighting matrix	.733	.888	43.2
	(.222)	(.017)	[.0000]
Eliminating outliers	.893	.890	41.4
	(.309)	(.024)	[.0000]
	Alternative Instrument Sets		
Debt–equity ratio	.345	.913	31.9
	(.268)	(.014)	[.0000]
Debt–equity ratio and spread	.739	.909	38.7
	(.220)	(.016)	[.0000]
Interest–Coverage Ratio	.157	.912	31.2
	(.139)	(.014)	[.0001]
Interest–Coverage Ratio (Positive Only)	.217	.917	23.2
	(.119)	(.015)	[.0016]
	(b) Informational Classes		
Group	.508	.793	11.9
	(.258)	(.055)	[.1561]
Independent	.671	.911	34.7
	(.237)	(.016)	[.0000]
Concentrated	.625	.814	15.4
	(.421)	(.030)	[.0516]
Dispersed	.678	.930	22.0
	(.228)	(.013)	[.0050]
Mature	.726	.893	20.4
	(.332)	(.018)	[.0048]
Young	.481	.923	25.7
	(.253)	(.018)	[.0012]

Note: Standard errors are in parentheses and p-values are in square brackets. The estimation method is GMM. A large J statistic (Hansen (1982)) implies a rejection of the symmetric information specification which is presented in equation (6). The instruments are a constant and the t-2 values of Y/K, C/K, p', S, τ, δ, I/K, and $(I/K)^2$, except in the portion of the table labeled "Alternative Instrument Sets," where S is replaced by the variable listed in the first column. In the last row of part (a), the sample is restricted to firms for which the interest coverage ratio is never negative. In part (b), classes of firms are arranged in contrasting pairs. We expect group, concentrated ownership, and mature firms to find it easier to credibly communicate private information.

which are less than one, are consistent with either perfect competition and mildly decreasing returns to scale, or imperfect competition and more strongly decreasing returns. The adjustment cost coefficient is positive and significant. The J-statistic for the symmetric information specification is 39.2, which implies a rejection at a marginal significance level of 0.0001. This rejection of the symmetric information specification parallels the findings of Gertler et al. (1991) and Whited (1992) on U.S. aggregate and firm level data, respectively.

We checked the robustness of the results in several ways. First, the results just presented are based on a risk-free interest rate (the 1 month T-bill rate) as our measure of r. It is possible that the symmetric information specification is rejected because the risk-free interest rate does not incorporate variations in risk. To test this, we use a corporate bond rate to construct the discount factor in row 2 of table 1.[6] The use of an interest rate that includes a risk premium makes very little difference in the J-statistic, which still rejects the symmetric information model at the 0.0001 level. Second, instead of using the weighting matrix for the symmetric information model, we construct the weighting matrix for the agency cost model and use it in estimating the symmetric information specification. The results, reported in the third row, show a J-statistic of 43.2. Third, we eliminate outlier firms (based on I/K).[7] The results, presented in the fourth row, show a J-statistic of 41.4.

We also checked the robustness of the results to the choice of instruments. The spread is included in our base instrument set because it may be linked to agency costs. However, it is possible that it is associated with other sources of misspecification. One way to check this is to use as instruments balance sheet variables which have been linked to finance constraints in previous work. See, for example, Whited (1992). We first replace the risky spread in the instrument set with the debt–equity ratio, and then use both the debt–equity ratio and the risky spread as instruments. Next, we consider the interest–coverage ratio in place of the debt–equity ratio. The interest–coverage ratio, defined as the ratio of interest payments to the sum of interest payments and cash flow, will typically be higher for a firm that has encountered negative shocks that reduce cash flow and/or increase interest payments. But if a firm faces sufficiently negative shocks, the ratio can become negative (e.g., if cash flow becomes negative). We therefore consider the full sample and the subset of firms for which the interest–coverage ratio is never negative. These results are reported in the middle portion of table 1. For all these instrument sets, the J test strongly rejects the symmetric information model.

Results for the agency cost model are reported in table 2. The estimates for ϕ and ψ are omitted to conserve space. The $\hat{\psi}$'s are similar to the ones in table 1, while the $\hat{\phi}$'s are smaller and less precisely estimated than in the symmetric information specification. The coefficient on the spread $\hat{\gamma}_1$, gives an indication of the importance of financing constraints. In the first row of table 2, the point estimate of γ_1 is 7.2 with a standard error of 1.6. This suggests that the impact of financing constraints is not only statistically significant but economically important. To give a sense of how large the effect is, we consider a one standard deviation increase in the spread. For our time sample, this is equal to an increase of 65 basis points. Such an increase in the spread has the same effect on the shadow cost of finance as a 468 basis point increase in real interest rates. This increase in shadow cost is roughly equal to the mean real interest rate of 4%. Gertler et al. (1991) and Whited (1992) also find evidence of a statistically significant and economically important wedge between the market interest rate and the shadow cost of finance.

A further and more formal test of the symmetric information model against the agency cost model is to construct

[6] We also considered a 30-day banker's acceptance as a measure of r. This yielded a J-statistic of 36.0 and similar parameter estimates.

[7] We define outliers as those firms with I/K negative or greater than two in any year. This removes seven firms from the sample. The estimates in rows three and four use the corporate bond rate to construct r. The results based on the risk-free rate were similar.

TABLE 2.—AGENCY COST SPECIFICATION

	γ_0	γ_1	γ_2	J	NW	Favored Specification
			(a) Full Panel			
Risk-free rate	−.112	7.2		14.4	25.3	Agency Cost
	(.075)	(1.6)		[.0255]	[.0000]	
Risky rate	−.152	8.6		10.5	32.7	Agency Cost
	(.007)	(1.7)		[.1069]	[.0000]	
Alternative weighting matrix	−.147	8.3		10.6	32.7	Agency Cost
	(.077)	(1.7)		[.0602]	[.0000]	
Removing outliers	−.085	7.7		18.4	15.4	Agency Cost
	(.085)	(2.6)		[.0025]	[.0005]	
Debt–equity ratio	−.209	10.2	.132	24.6	11.0	Agency Cost
	(.146)	(2.3)	(.061)	[.0002]	[.0039]	
Interest–coverage ratio	−.118	14.6	−.098	19.1	26.1	Agency Cost
	(.118)	(2.5)	(.056)	[.0018]	[.0000]	
			(b) Informational Classes			
Group	.139	2.7		2.3	12.3	Agency Cost
	(.101)	(2.0)		[.8927]	[.0021]	
	.143	2.6	.001	2.52	12.3	Agency Cost
	(.153)	(2.4)	(.037)	[.7739]	[.0021]	
Independent	.158	8.0		17.5	19.4	Agency Cost
	(.095)	(2.0)		[.0075]	[.0000]	
	−.175	10.1	.126	23.0	4.59	Agency Cost
	(.152)	(2.4)	(.060)	[.0003]	[.1007]	
Concentrated	.020	9.3		23.0	6.3	Symmetric Info
	(.085)	(2.3)		[.0008]	[.0429]	
	.179	8.0	−.053	18.3	.246	Agency Cost
	(.099)	(2.5)	(.017)	[.0025]	[.8841]	
Dispersed	−.126	7.5		11.2	11.4	Agency Cost
	(.125)	(2.5)		[.0816]	[.0034]	
	−.162	11.1	.092	18.3	.142	Agency Cost
	(.164)	(3.1)	(.054)	[.0026]	[.9312]	
Mature	.048	10.8		23.6	2.2	Symmetric Info
	(.092)	(2.4)		[.0003]	[.3381]	
	.257	9.8	−.077	16.1	4.75	Agency Cost
	(.127)	(2.4)	(.028)	[.0067]	[.0929]	
Young	.007	5.5		12.9	12.4	Symmetric Info
	(.121)	(2.4)		[.0452]	[.0020]	
	−.093	9.8	.091	10.8	6.86	Agency Cost
	(.162)	(3.0)	(.052)	[.0564]	[.0323]	

Note: Standard errors are in parentheses and p-values are in square brackets. γ_1 and γ_2 measure the effect of shocks to internal net worth on the shadow cost of finance. The *NW* statistic (Newey and West (1987b)) compares the symmetric information specification (equation (6)) with the agency cost specification (equation (12)). The favoured specification is listed in the last column. See the notes to table 1 for other details of estimation and testing.

the χ^2 statistic suggested by Newey and West (1987b). The intuition for the test is that if a model is incorrectly specified, the J-statistic for the misspecified model will tend to be large. The difference in J-statistics between two models holding the weighting matrix fixed provides a test of whether the improvement in specification is statistically significant. This difference of two J-statistics is distributed as χ^2 with degrees of freedom equal to the number of omitted parameters. The unrestricted model in this context is the agency cost specification since it places no restrictions on the γ parameters. Its weighting matrix is therefore used to construct the *NW* tests. The symmetric information model involves two fewer parameters, so all the statistics labelled *NW* in the tables have two degrees of freedom. The *NW*-statistic is 25.3, which is significant at the 0.0001 level, showing an improvement moving from the symmetric information to the agency cost model.

To check the robustness of the *NW* tests, we use a risky interest rate to construct r. When risk is incorporated into the discount rate, the *NW*-statistic continues to reject the symmetric information specification in favour of the agency cost specification. In addition, the estimate of γ_1 in the agency cost model increases slightly from 7.2 to 8.6. Using the weighting matrix from the symmetric information specification to estimate the agency cost model and eliminating outliers from the sample have no substantial effect on the results, as shown in rows three and four of table 2. We then add balance sheet variables to the parameterization of the shadow cost of financing constraints. In row five when we add the debt–equity ratio, the estimate of γ_1 rises. Hence the inclusion of firm-specific measures does not seem to reduce the importance of aggregate shocks to net worth, as summarized by the risky-spread. There is also evidence that firm-specific shocks matter since $\hat{\gamma}_2$ is significantly different from zero. The point estimate of γ_2 implies that a one standard deviation shock to the debt–equity ratio would increase the shadow cost of finance by 78 basis points. The Newey-West test again strongly rejects the symmetric information model in favour of the agency cost specification. The results in row six for the interest–coverage ratio are similar, except that γ_2 is imprecisely estimated.

B. *Identifying Firms in Different Informational Positions*

A distinctive feature of our work is that we identify firms which face different degrees of informational asymmetry with respect to sources of financing and test whether their investment behavior differs in the way predicted by the agency cost interpretation. To do this, we have gathered additional information about the firms in our sample which is typically not available in the firms' financial statements. We have then classified firms in three different ways which are intended to capture the firms' informational positions.

One way to mitigate informational problems is through groupings of firms. By entering into long-term relationships, firms introduce reputational effects which allow members to credibly communicate private information about the quality of individual projects to other members of the group. Hoshi et al. (1991) have previously used a similar approach to classify firms. They find that firms which are members of Japanese groups (keiritsu) show less sensitivity of investment to cash flow than independent firms. While Canada does not have a direct analog of the Japanese keiritsu, there are several major Canadian industrial groups which are closer in some respects to the Japanese groups than to organizations in the United States. Many Canadian enterprises are organized as groups, many of which have been associated with a particular family or individual (e.g., the Belzberg, Black, Bronfman, Desmarais, and Reichman groups), although there are exceptions (e.g., the Canadian Pacific group).

Data on firm affiliations were obtained from *Intercorporate Ownership 1984* (published by Statistics Canada), which contains a list of the firms associated with groups and the nature of the connection, i.e., whether it is based on share ownership or some other relationship. We classify a firm as a group member if the firm is majority-owned or effectively controlled by one of the Canadian industrial groups, or if one of the Canadian industrial groups is the largest shareholder. The remaining firms are classified as independent. Regression results which test if the symmetric information model holds for group and independent firms respectively are presented in panel (b) of table 1. There is no evidence that the investment behavior of group firms violates the symmetric information hypothesis; the marginal significance level of the J-test is 0.1561. There is strong evidence, however, that the behavior of independent firms is inconsistent with symmetric information. The J-test is significant at the 0.0001 level.

The second classification we examine separates firms with concentrated ownership from those with dispersed ownership. Data on ownership concentration were also obtained from Statistics Canada's *Intercorporate Ownership 1984*. A firm is classified as concentrated if either one shareholder holds 50% or more of the shares, or the firm is effectively controlled by another firm.[8] The remaining firms are classified as having dispersed ownership. The more concentrated the ownership of the firm, the more closely managers' interests should coincide with those of shareholders since the shareholders' free-rider problem associated with monitoring management performance is reduced. Thus the more concentrated the ownership of the firm, the smaller the risk faced by a potential investor or lender that the firm will misrepresent the quality of its investment project, and thus the smaller the agency problem. As table 1(b) shows, the data fail to reject the symmetric information specification for firms with concentrated ownership. The symmetric information specification is strongly rejected for firms with dispersed ownership.

The third classification separates mature firms from young firms. Mature firms are less likely to face informational problems, both because lenders will tend to know more about firms that have been visible for an extended period of time, and because mature firms can credibly enter into repeated relationships with lenders. We classify those firms which have been tracked by Laval (the Canadian equivalent of CRSP) since its inception in 1963 as mature; the remaining firms are classified as young.[9] Separate estimates for mature and young firms are also presented in table 1(b). This is probably the weakest of our classifications because all of the firms classified as ''young'' will have been in existence for at least thirteen years by the end of our sample period. Nonetheless, there is some evidence of a difference between the firms we classify as young and mature. As predicted by the agency cost model, the J-test rejects the symmetric information specification more strongly for young firms than for mature firms.

It is of interest to note that the characteristics that we consider are not highly correlated. The correlation between mature firms and those with concentrated ownership is 0.16, between mature firms and members of an industrial group is 0.23, and between group members and firms with concentrated ownership is 0.27. Moreover, the classes do not seem to be simply proxies for the size of the firm. We divided the firms in our sample into two equal classes based on their size as measured by the market value of equity in 1973. The J-statistic was actually slightly greater for large firms but the difference was negligible, with the marginal significance level being 0.0001 for large firms and 0.0006 for small firms.

To assess the sensitivity of the results to the choice of instruments, we replace the risky spread by either the debt–equity ratio or the interest–coverage ratio in the instrument set. In the majority of cases, the data fail to reject the symmetric information model for group firms, firms with concentrated ownership, and mature firms. In all cases, the

[8] For example, Brascan is effectively controlled by the Bronfman family, although they only own 43% of the shares.

[9] Firms for which this information is not available have been retained in the sample but excluded from comparisons of young and mature firms.

symmetric information model is rejected for independent firms, firms with dispersed ownership, and young firms. Thus, tests of overidentifying restrictions tend not to reject the symmetric information model for firms which are in a strong informational position with respect to capital markets. The symmetric information model is rejected for each class of firms which is in a weak informational position. More generally, the marginal significance levels are lower for firms which are in a weak informational position.

Panel (b) of table 2 provides disaggregated estimates of the agency cost specification. We first focus on estimations which use only the risky spread to parameterize the Lagrange multiplier. For most classes of firms, the magnitude of $\hat{\gamma}_1$ is greater than 5.0 and the t-statistic is more than 2. The exception is group firms, for which the point estimate of γ_1 is 2.7 and is insignificantly different from 0. Thus, for group firms, the results show no significant evidence of a wedge between the shadow cost of finance and the market interest rate. But for independent firms, shocks to net worth can cause a divergence between the market interest rate and the shadow cost of finance. Table 2(b) also indicates that for all three classes of firms in a weak informational position (independent, dispersed ownership, and young), the *NW* test is highly significant and favors the agency cost specification. For two of the three classes of firms in a strong informational position, the *NW*-statistic favors the symmetric information specification.

Disaggregated estimates for the agency cost model using both the risky spread and the debt–equity ratio to parameterize the Lagrange multiplier are also given in table 2(b). For group firms, firms with concentrated ownership, and mature firms, the estimate of γ_1 falls relative to the parameterization without the debt–equity ratio. For independent firms, firms with dispersed ownership, and young firms, the opposite is true. For group firms, $\hat{\gamma}_2$ is approximately zero, implying an insensitivity of the shadow cost of finance to the debt–equity ratio. For independent firms, $\hat{\gamma}_2$ is positive and statistically significant. This provides evidence that firm-specific shocks to net worth are important even after controlling for aggregate shocks. For all classes of firms, the agency cost model is preferred over the symmetric information model.

We took several steps to check the robustness of the results. First, we examined whether the results are sensitive to whether the firm has public debt outstanding. Results from estimations which exclude firms that do not have public debt are similar. Second, we estimated a specification using the interest–coverage ratio instead of the debt–equity ratio as X_{t-1}. Many of the patterns are similar. The estimates for γ_1 are small and insignificantly different from zero for group firms, but large and highly significant for independent firms. Relative to the results in table 2(b) for the parameterization without the debt–equity ratio, $\hat{\gamma}_1$ rises for all classes in a weak informational position, but falls for all classes in a strong informational position. The agency cost model is preferred for all classes of firms. One difference is that $\hat{\gamma}_2$ is less precisely estimated as all the t-statistics are less than two.

IV. Conclusion

Using Canadian panel data, we compare a symmetric information specification of investment to one in which agency costs play an important role. For the full panel of 199 firms, we strongly reject the symmetric information specification. Results for the full panel are more supportive of an agency cost specification in which firms sometimes face finance constraints. Coefficient estimates suggest that shocks to net worth, as reflected in the risky spread and firm-specific balance sheet variables, have economically important and statistically significant effects on the shadow cost of finance.

Using distinctive Canadian institutional features to classify firms, we find important differences between firms which are in a strong position to credibly communicate private information and firms in a weak informational position. The symmetric information specification is always more strongly rejected by firms in a weak informational position. For these firms, the agency cost specification is favored.

DATA APPENDIX

Firm-specific depreciation rates are constructed from firms' reported depreciation based on the procedure discussed in Salinger and Summers (1983). A recursive formula is used to calculate the capital stock, which evolves as $K_t = K_{t-1}p^I_{t-1}/p^I_t(1 - \delta) + I_t$, where p^I_t is the relative price of investment, the implicit price index for business investment in machinery and equipment (CANSIM series D11123) relative to the implicit price index for final domestic demand (CANSIM series D11130).

We use the difference in the annual averages of the 30-day bankers' acceptance rate and the 30-day T-bill rate as our base measure of the risky spread. The quantity Y in the Euler equations is substituted by the value of sales from the firm's balance sheets. The balance sheets do not directly report variable costs, but they do report operating income, which is revenue less variable costs. We can therefore construct variable costs by subtracting operating income from revenue. The investment data also come directly from the firm's balance sheets. For τ we use the statutory corporate tax rate; since Canada has a separate corporate tax rate for manufacturing and mining, we choose the appropriate tax rate for each firm, which means that there is some cross-sectional variation in tax rates. We use the statutory investment tax credit rate. The present value of depreciation allowances for tax purposes is calculated using the method suggested by Salinger and Summers (1983), a method which has been widely used in investment studies on panel data.

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