

# The Corporate Propensity to Save

Leigh A. Riddick  
American University

Toni M. Whited\*  
University of Wisconsin, Madison

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\*Corresponding author. School of Business, University of Wisconsin, Madison, WI 53706-1323. 608-262-6508. [twhited@bus.wisc.edu](mailto:twhited@bus.wisc.edu). We are grateful for helpful comments and suggestions from Tor-Erik Bakke, Morris Davis, Mark Garmaise, Robert Hauswald, Chris Hennessy, Holger Mueller, Michael Roberts and seminar participants at the American University, Duke, HEC-Lausanne, HEC-Paris, Northwestern, and the University of Wisconsin, Madison. Bobby Hart of Thompson Financial was helpful in data acquisition, and Michael R. Sullivan provided research assistance. Leigh Riddick also wishes to acknowledge research support from the Kogod School of Business and American University.

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## Abstract

We study corporate saving in a stochastic, dynamic model of the firm with endogenous choices of external finance, distributions, cash, and investment. Intertemporal trade-offs between costly external finance and interest income taxation determine optimal savings. Unlike static models, our model produces negative propensities to save out of income because firms lower cash reserves to invest after receiving good productivity news, and vice versa. OLS regressions using international data replicate positive saving propensities found previously. However, treating measurement error in Tobin's  $q$  produces negative propensities, consistent with our model. Empirically, income uncertainty matters more for saving than the cost of external finance.

How does the cost of external finance affect corporate decisions—both real and financial? Whatever the answer, it must incorporate the idea that one cannot understand real decisions without understanding how they interact with financial decisions, and vice versa. We apply this general principle to ask what factors affect corporate saving policy, that is, the propensity for a firm to funnel its cash flow into its stock of cash holdings. In so doing, we draw a distinction between why firms hold cash and why firms change their given cash holdings, that is, save or dissave.

We develop a model that provides several surprising insights. In our infinite-horizon framework firms invest, save, produce, and make distributions in the face of uncertainty, physical adjustment costs, taxation, and costly external finance. We solve the model numerically, using the solution to generate an exact testable implication concerning a commonly used measure of the corporate propensity to save—the sign of the coefficient on cash flow in a linear regression of the change in cash levels (saving) on Tobin’s  $q$  and cash flow. Most model simulations produce a negative coefficient on cash flow, which becomes increasingly negative as external finance becomes more costly, as the serial correlation in income rises, and as the degree of income uncertainty falls. We then estimate this saving regression using international firm-level data, and our data analysis supports the model’s predictions.

Although the result of a negative sensitivity of saving to cash flow is, at first, counterintuitive, the intuition behind the model’s results is straightforward. Because interest on any cash balances is taxed, the firm faces a dynamic trade-off between this tax penalty and the reduction in expected future financing costs conferred by holding cash. The firm’s optimal saving policy therefore depends not only on the cost of external finance, but also on its expected future financing needs, which, in turn, depend on its technology and especially on the nature of the uncertainty it faces in its income.

For example, if the firm faces positively serially correlated income, a good shock means that the firm’s capital goods will be productive in the future, and it invests. It also knows that it will be unlikely to need external finance in the future. These two effects lead to a propensity to dissave out of cash flow. On the other hand, we find the opposite result if the firm faces highly negatively correlated income shocks. In this case, for example, if a firm experiences a good shock, it knows that it will likely experience a bad shock in the future. It therefore expects the productivity of

capital goods to decline, which leads to a decrease in investment. The firm also anticipates an income shortfall in the future, which may force the firm to find outside finance for any desired future investment projects. These two effects cause an increase in saving. On a priori grounds this second result is unlikely inasmuch as most firms have positively serially correlated cash flows.

The negative relation between saving and cash flow becomes stronger as the cost of external finance rises because the firm's optimal *level* of cash increases with the cost of external finance. In comparison with a low-cost firm, a high-cost firm therefore has more slack with which to respond to profit shocks, and it saves or dissaves more aggressively in large part to counteract these shocks. Although the levels and changes in cash for a firm are clearly related, a high cash level does not necessarily imply a high, positive sensitivity of the change in cash (saving) to cash flow; nor does a low cash level imply a low sensitivity. This distinction between levels and sensitivities would be impossible to uncover, for example, in a static, two-period model because in such a setting the change in cash is indistinguishable from the level. A dynamic model such as ours is therefore essential to understanding the trade-offs that affect corporate saving.

We test our model's prediction that the cash-flow coefficient is negative in a regression of corporate saving on Tobin's  $q$  and cash flow, using firm-level data from six countries—Canada, France, Germany, Japan, the United Kingdom, and the United States. In support of the findings in Almeida, Campello, and Weisbach (2004), OLS regressions produce a positive coefficient on cash flow for all six countries. We also confirm their result that this coefficient is higher for small than for large firms. However, when we correct for the well-documented measurement error in  $q$ , following Erickson and Whited (2000, 2002), we find negative coefficients on cash flow in all six countries, and many are significant. Firms with highly positively serially correlated income also have significantly larger negative saving estimated propensities than firms with serially uncorrelated income. Further, we find a negative relation between saving propensities and the variability of income. Because firms typically categorized as facing finance constraints also have highly variable income shocks, their cash-flow coefficients are closer to zero than those of their unconstrained counterparts. In sum, from our simulations and empirical evidence we conclude that income shocks are at least as important as the cost of external finance in determining corporate saving. Accordingly, saving propensities

cannot be used as summary measures of the cost of external finance.

Our results may seem somewhat puzzling in light of the argument that measurement error in Tobin's  $q$  should not matter in a regression of the change in cash levels (rather than investment) on Tobin's  $q$  and cash flow. As argued in Almeida, et al. (2004), cash flow contains information about future investment opportunities, and this information leads to the significance of cash flow in an investment, but not in a saving, regression. Implicit in this argument is the idea that cash flow contains little information about future saving opportunities. If this argument is correct, then using a measurement-error consistent estimator should confer no advantage over OLS. As explained in Erickson and Whited (2000), however, measurement error in Tobin's  $q$  can bias any regression coefficient as long as the regressor in question is correlated with true, unobserved  $q$ . The information about future investment opportunities in cash flow leads naturally to a positive correlation between  $q$  and cash flow—a correlation that always biases the cash-flow coefficient in the presence of a noisy measure of Tobin's  $q$ , regardless of the left-hand-side variable in the regression. We are able to avoid this issue here by correcting for measurement error in  $q$ .

Because one of the main contributions of our paper is a richer framework for investigating corporate saving, it is important to identify the advantages of our approach in comparison with what has been the workhorse of the empirical corporate cash literature: testing empirical predictions that take the form of the sign of a partial derivative derived from a static model. These predictions by nature hold all other relevant variables constant, and they typically give little guidance concerning the construction of a regression that does not suffer from endogeneity. In contrast, our model simulations provide an exact regression specification that accounts for several decisions that are made jointly. Moreover, because our empirical predictions take the form of linear projections, they do not, by definition, suffer from endogeneity.

Using a dynamic model to generate a regression has precedents in Whited (2006) and Caggese (2006), and it is closely related to calibration inasmuch as both approaches use carefully parameterized models to explain observed data moments. What we do differs from calibration in that we place a greater emphasis on statistical inference. Our approach is also related to the simulated moments estimation in Hennessy and Whited (2005, 2006). On one hand, estimation via simulated

moments is preferable because it provides a direct link between the underlying theory and its tests. In contrast, our approach only provides an indirect link between theory and tests. Although we derive the exact form of a regression to estimate, we do not link regression coefficients directly to structural parameters. On the other hand, structural estimation faces the issue that any inference drawn could be fragile with respect to the assumptions used to estimate the model. Because our approach does not require that our model be exactly true with regard to our data moments of interest, it is more robust to problems of misspecification.

The theoretical model in this paper is most closely related to that in Whited (2006), in which a firm chooses investment and cash holdings in the face of costly external equity finance and fixed costs of capital adjustment. We extend the model by including a corporate income tax and convex adjustment costs, and we examine empirically the model's implications for saving rather than for investment, as in Whited (2006). Our model is also closely related to the one in Eisfeldt and Rampini (2006), which characterizes the business-cycle properties of aggregate liquidity. They calibrate a general-equilibrium model with a rich specification of uncertainty. Although many of the same economic mechanisms at work in their model also operate in ours, the focus of the two papers is quite different in that we are interested in directly testing the implications of the model at the firm level, instead of calibration at the aggregate level. Another closely related paper is Gamba and Triantis (2006). Their model is quite general, allowing for cash holding as well as separate debt and equity finance, although, unlike us, they omit physical adjustment costs. Their main contribution is an explanation of how debt flotation costs can lead to simultaneous cash and debt holdings. Once again, however, their paper is entirely theoretical.

Our model is also related to those recent theoretical and empirical papers that use dynamic frameworks to tackle other corporate finance questions. Theoretical papers include Gomes (2001), who studies the sensitivity of investment to cash flow in a general equilibrium model; Cooley and Quadrini (2001), who study entry and exit in an industry-equilibrium model; and Moyen (2004), who studies investment in a partial-equilibrium model of financially constrained and unconstrained firms. Empirical papers include Cooper and Ejarque (2003), who estimate a partial equilibrium model to understand investment-cash flow sensitivities; Caggese (2006), who examines the effects

of finance constraints on variable capital; and Hennessy and Whited (2005, 2006), who estimate dynamic models of corporate capital structure and the cost of external finance.

Our empirical work is closely related to Almeida, et al. (2004), who investigate firms' tendencies to save in the face of costly external finance; Khurana, Pereira, and Martin (2006), who replicate the results in Almeida, et al. (2004) on data from several countries; Acharya, Almeida, and Campello (2006), who examine both the propensity to save out of cash flow and the propensity to issue debt; and Sufi (2006), who uses saving-cash flow sensitivities as an explicit metric for gauging the severity of the cost of external finance. In contrast to the studies in this group we motivate our tests with a dynamic model and use measurement-error consistent estimators. These features of our research design lead us to very different conclusions.

Our work is also tangentially related to recent empirical work on the determinants of the level of (as opposed to the change in) corporate cash holdings, such as Kim, Mauer, and Sherman (1998), Opler, Pinkowitz, Stulz, and Williamson (1999), Pinkowitz and Williamson (2001), Dittmar and Mahrt-Smith (2005), Faulkender and Wang (2006), and Foley, Hartzell, Titman, and Twite (2006).

The paper is organized as follows. Section 1 presents the model. Section 2 describes the model simulation and its results. Section 3 describes the data, Section 4 presents the estimation procedure and results, and Section 5 concludes. The Appendix contains details regarding the estimation and a Monte Carlo simulation designed to determine the estimators' finite-sample performance.

## 1. A Model of Cash Holding

To motivate our empirical work, we consider a discrete-time, infinite-horizon, partial-equilibrium model of investment and saving. First we describe the technology, taxation, and financial frictions. Then we move on to a description of the optimal financing policies.

### 1.1. Technology and Financing

A risk-neutral producer uses capital,  $k$ , to produce output. The producer's per period profit function is given by  $\pi(k, z)$ , in which  $\pi(k, z)$  is continuous,  $\pi(0, z) = 0$ ,  $\pi_z(k, z) > 0$ ,  $\pi_k(k, z) > 0$ ,  $\pi_{kk}(k, z) < 0$ , and  $\lim_{k \rightarrow \infty} \pi_k(k, z) = 0$ . The profit function can be thought of as a reduced-form production function in which variable factors have already been maximized out of the problem.

Concavity of  $\pi(k, z)$  results from decreasing returns in production, a downward sloping demand curve, or both. The combination demand and productivity shock is denoted by  $z$ . It is observed by the producer before he makes his current period decisions. The shock  $z$  takes values in  $[\underline{z}, \bar{z}]$  and follows a first-order Markov process with transition probability  $g(z', z)$ , in which a prime indicates a variable in the next period;  $g(z', z)$  has the Feller property.

We assume that corporate income is taxed at a rate  $\tau_c$ . For simplicity, we do not model personal interest and dividend taxes. What is important for our model is the existence of a tax penalty for saving, which is consistent with recent U.S. tax code. See Hennessy and Whited (2005).

Without loss of generality,  $k$  lies in a compact set. As in Gomes (2001), define  $\bar{k}$  as

$$(1 - \tau_c) \pi(\bar{k}, \bar{z}) - d\bar{k} \equiv 0, \quad (1)$$

in which  $d$  is the constant rate of depreciation,  $0 < d < 1$ . Concavity of  $\pi(k, z)$  and  $\lim_{k \rightarrow \infty} \pi_k(k, z) = 0$  ensure that  $\bar{k}$  is a well-defined quantity. Because  $k > \bar{k}$  is not economically profitable,  $k$  lies in the interval  $[0, \bar{k}]$ . Compactness of the state space and continuity of  $\pi(k, z)$  ensure that  $\pi(k, z)$  is bounded. Investment,  $I$ , is defined as

$$I \equiv k' - (1 - d)k. \quad (2)$$

The firm faces both real and financial frictions. First, the firm purchases and sells capital at a price of 1. When it does so, it incurs adjustment costs according to

$$A(k, k') = ck\Phi_i + \frac{a}{2} \left( \frac{k'}{k} - (1 - d) \right)^2 k. \quad (3)$$

The functional form of (3) encompasses both fixed and smooth adjustment costs. The fixed component is captured by the first term,  $ck\Phi_i$ , in which  $c$  is a constant, and  $\Phi_i$  is an indicator equal to 1 if investment is nonzero, and 0 otherwise. The fixed cost is proportional to the capital stock so that the firm has no incentive to grow out of the fixed cost.<sup>1</sup> The second term is a standard quadratic adjustment cost function commonly found in the empirical investment literature. Although curvature of the profit function acts to smooth investment over time in the same way that the quadratic component of (3) does, we include the quadratic component to isolate the effects of

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<sup>1</sup>Replacing  $ck$  with a fixed number,  $F$ , changes the analysis little because the capital stock is bounded.

smooth adjustment costs. In contrast, curvature of the profit function not only affects investment smoothing but also the relation between firm value and profit.

We now discuss financing. The firm can hold cash balances,  $p$ , via a riskless one-period discount bond that earns an interest rate  $r$ . To ensure compactness of the choice set, we assume an arbitrarily high upper bound on corporate saving,  $\bar{p}$ . This upper bound is imposed without loss of generality, because our taxation assumptions ensure bounded saving. All external finance takes the form of equity. Although it would be inappropriate for the study of capital structure, this simplification allows us to highlight the interaction between technology, finance constraints, and cash holdings. Further, the simple structure does not affect the qualitative outcome of the simulations that follow. To preserve tractability, we do not model costs of external equity as the outcome of an asymmetric information problem. Instead, we capture adverse selection costs and underwriting fees in a reduced-form fashion. Accordingly, we define the excess of cash inflows over cash outflows as

$$e(k, k', p, p', z) \equiv (1 - \tau_c) \pi(k, z) + p - \frac{p'}{(1 + r(1 - \tau_c))} - (k' - (1 - d)k) - A(k, k'). \quad (4)$$

If  $e(k, k', p, p', z) > 0$ , the firm is making distributions to shareholders, and if  $e(k, k', p, p', z) < 0$ , the firm is issuing equity. The external equity-cost function is linear-quadratic and weakly convex:

$$\begin{aligned} \phi(e(k, k', p, p', z)) &\equiv \Phi_e \left( -\lambda_0 + \lambda_1 e(k, p, k', p', z) - \frac{1}{2} \lambda_2 e(k, p, k', p', z)^2 \right) \\ \lambda_i &\geq 0 \quad i = 0, 1, 2, \end{aligned}$$

in which  $\Phi_e$  is an indicator equal to 1 if  $e(k, p, k', p', z) < 0$ , and 0 otherwise. Convexity of  $\phi(e(k, p, k', p', z))$  is consistent with the evidence on underwriting fees in Altinkilic and Hansen (2000).

The firm chooses  $(k', p')$  each period to maximize the value of expected future cash flows, discounting at the opportunity cost of funds,  $r$ . The Bellman equation for the problem is

$$V(k, p, z) = \max_{k', p'} \left\{ e(k, k', p, p', z) + \phi(e(k, k', p, p', z)) + \frac{1}{1 + r} \int V(k', p', z') dg(z', z) \right\}. \quad (5)$$

The model satisfies the conditions for Theorem 9.6 in Stokey and Lucas (1989), which guarantees a solution for (6) and (7). Theorem 9.8 in Stokey and Lucas (1989) ensures a unique optimal policy function,  $\{k', p'\} = h(k, p, z)$ , if  $e(k, k', p, p', z) + \phi(e(k, k', p, p', z))$  is weakly concave in its first

and third arguments. This requirement puts easily verified restrictions on  $\phi(\cdot)$ , which are satisfied by all functional forms chosen below.

## 1.2. Optimal Financial Policies

This subsection develops the intuition behind the model by examining its optimality conditions. To simplify the exposition of optimal policies, we assume in this subsection that  $V$  is concave and once differentiable. These assumptions are not necessary for the existence of a solution to (5) or of an optimal policy function. We present optimal financial policies, heuristically, in two steps. First, we determine optimal financing under the assumption that the manager ignores the fixed costs of external equity; that is, he or she treats  $\lambda_0 = 0$ . Second, we determine whether the intra-marginal benefits of equity issuance justify the fixed cost.

The optimal interior financial policy, obtained by solving the optimization problem (5), satisfies

$$1 + (\lambda_1 - \lambda_2 e) \Phi_e = \frac{1 + r(1 - \tau_c)}{1 + r} \int V_2(k', p', z') dg(z', z). \quad (6)$$

The right side represents the shadow value of cash balances, and the left side represents the marginal cost of external equity finance. To develop the intuition behind the optimal policy, we use the envelope condition to rewrite (6) as:

$$1 + (\lambda_1 - \lambda_2 e) \Phi_e = \frac{1 + r(1 - \tau_c)}{1 + r} \int (1 + (\lambda_1 - \lambda_2 e') \Phi'_e) dg(z', z). \quad (7)$$

Rewriting (6) as (7) makes it clear that without costly external finance, equation (7) holds as an inequality. In this case the tax penalty for saving implies that the firm never saves; i.e.,  $p = p' = 0$ . In contrast, in the face of costly external finance, if a firm saves a dollar today, it reduces the probability of having to issue new equity tomorrow. It continues to save just until the gain from reducing future equity costs outweighs the tax penalty on saving.

In some instances the fixed costs of external equity will be larger than the intra-marginal gains from equity issuance. In these cases the firm is in a region of financial inertia in which it neither issues equity nor distributes funds to shareholders. Internal funds are the marginal source of funds and the firm saves any excess cash flows not used for positive NPV projects.

## 2. Simulations

We solve the model numerically and investigate its implications for reduced-form regressions via simulation. We first describe the parameterization of our baseline simulation. We then explain the experiments we perform on the model and the results of these experiments. We conclude by considering the empirical predictions given by the model and by examining the robustness of the model to our various simplifying assumptions.

### 2.1. Model Calibration

The profit function is given by

$$\pi(k, z) = zk^\theta, \tag{8}$$

We calibrate the value of  $\theta$  from the estimates of labor shares and mark-ups in Rotemberg and Woodford (1992, 1999). Their estimates, along with the assumption of a Cobb-Douglas production function and a constant-elasticity demand function, imply that  $\theta \approx 0.75$ .

To find values for the adjustment cost parameters,  $c$  and  $a$ , we turn to Cooper and Haltiwanger (2006), who find that both convex and fixed costs of adjustment affect investment. From their estimates we set  $c = 0.039$  and  $a = 0.049$ . We set the depreciation rate equal to 0.15, a figure approximately equal to the average in our data of the ratio of depreciation to the net capital stock.

Next, we specify a stochastic process for the shock  $z$ . Following Gomes (2001), we assume that  $z$  follows an  $AR(1)$  in logs,

$$\ln(z') = \rho \ln(z) + v', \tag{9}$$

in which  $v' \sim N(0, \sigma_v^2)$ . Our baseline parameter choices for  $\rho$  and  $\sigma_v$  are the averages of the estimates of these two parameters in Hennessy and Whited (2006):  $\rho = 0.66$  and  $\sigma_v = 0.121$ .

We again follow Hennessy and Whited (2006) to parameterize the financing function, setting  $\lambda_0 = 0.389$ ,  $\lambda_1 = 0.053$ , and  $\lambda_2 = 0.0002$ . These settings are from their estimates of the costs of external equity finance for large firms and are therefore conservative, lying only slightly above the figures for underwriting costs in Altinkilic and Hansen (2000). We set the interest rate,  $r$ , equal to 4%, which lies between the values chosen by Hennessy and Whited (2006) and Gomes (2001).

Finally, to find a numerical solution we need to specify a finite state space for the three state variables. We let the capital stock lie on the points

$$\left[\bar{k}(1-d)^{40}, \dots, \bar{k}(1-d), \bar{k}(1-d)^{1/2}, \bar{k}\right].$$

We let the productivity shock have 20 points of support, transforming (9) into a discrete-state Markov chain using the method in Tauchen (1986). We let  $p$  have 20 equally spaced points of support in the interval  $[0, \bar{p}]$ , in which  $\bar{p}$  is set to  $\bar{k}/2$ . This upper bound never binds for the optimal choices of  $p$  determined by the simulation procedure.

We solve the model via iteration on the Bellman equation, which produces the value function  $V(k, p, z)$  and the policy function  $\{k', p'\} = h(k, p, z)$ . In the subsequent model simulation, the space for  $z$  is expanded to include 80 points, with interpolation used to find corresponding values of  $V$ ,  $k$ , and  $p$ . The model simulation proceeds by taking a random draw from distribution of  $z'$  (conditional on  $z$ ), and then computing  $V(k', p', z')$  and  $h(k', p', z')$ . We use these computations to generate an artificial panel of firms by simulating the model for 10,000 identical firms over 200 time periods, keeping only the last 20 observations for each firm.

## 2.2. Experiments

We investigate two ways in which the model’s parameters affect the firm’s cash and saving policies. We first consider how the parameters affect the level of cash as a fraction of assets, which is defined in our model as the average of  $p/k$  over all of the observations in the simulated panel. We then examine how the parameters affect a measure of the change in cash (saving) that first appears in Almeida, et al. (2004). Dubbed “the cash-flow sensitivity of cash,” this measure is defined in our model as the regression coefficient,  $\alpha_1$ , in the following regression:

$$\frac{p' - p}{k} = \alpha_0 + \beta \frac{V(k, p, z)}{k} + \alpha_1 \frac{\pi(k, z)}{k} + u, \quad (10)$$

in which  $\alpha_0$ ,  $\alpha_1$ , and  $\beta$  are regression coefficients and  $u$  is a regression disturbance, which in our simulations is, by definition, orthogonal to the regressors.<sup>2</sup> This regression is estimated with all of the observations in the simulated panel. In thinking about the results that follow, it is crucial to

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<sup>2</sup>It is also possible to deflate the variables in (10) by total assets,  $(k + p)$ , rather than by  $k$ . Doing so, however, does not alter the qualitative results.

separate cash levels from the cash-flow sensitivity of cash, the latter of which we also refer to as the saving sensitivity or as the propensity to save.

We examine the sensitivity of these two gauges of cash policy to eight key model parameters: the variance and serial correlation of income; the three equity-cost parameters,  $\lambda_0$ ,  $\lambda_1$ , and  $\lambda_2$ ; the curvature of the profit function,  $\theta$ ; and the fixed and quadratic adjustment cost parameters,  $c$  and  $a$ . In each of the experiments that follow, we set all but one of the parameters equal to their baseline levels as defined above, allowing the free parameter to range within a given interval. We allow  $\theta$  to range between 0.6 and 0.9,  $\rho$  to range from -0.8 to 0.8,  $\sigma_v$  to range from 0.075 to 0.15,  $\lambda_0$  to range from 0 to 0.8,  $\lambda_1$  to range from 0 to 0.1,  $\lambda_2$  to range from 0 to 0.0004,  $c$  to range from 0 to 0.8, and  $a$  to range from 0 to 0.1.

Figure 1 illustrates the dependence on the model parameters of our first measure—the ratio of the level of cash holdings to the capital stock. We first examine the parameters that govern the stochastic shock process. The first panel shows that the relation between the serial correlation of income,  $\rho$ , and cash holdings is slightly u-shaped. For both highly positively and highly negatively correlated shocks, the firm holds high cash balances, choosing lower balances if the shocks are less highly correlated. Two separate effects explain this result. First, as the serial correlation of the income process increases, the firm tends to invest in larger amounts because a positive productivity shock signals not only that capital is productive today, but also that it will continue to be productive in the future. The firm therefore wants higher cash balances so that it will be less likely to need external finance when it makes these large investments. Second, the higher the serial correlation of an  $AR(1)$  process, the higher its variance. If the firm faces an uncertain environment, it expects to tap external finance more often, and it holds higher cash balances. Both effects operate in the same direction for high levels of  $\rho$ , but they appear to offset each other for levels near zero. For levels of  $\rho$  far below zero, the second effect dominates.<sup>3</sup> The intuition about the variance of the process is also evident in the second panel, which depicts a positive relation between cash holdings and  $\sigma_v$ , the variance of the innovations to  $z$ .

We next examine the effects of the cost of external finance. The third through fifth panels

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<sup>3</sup>Simulations in which  $\rho$  increases but the variance of the process is held constant produces a similar result, except that the slight rise in cash holdings for very low  $\rho$  flattens out.

illustrate the relations between each of the external finance parameters and cash holdings. Not surprisingly, the third and fourth panels show that cash increases with the fixed and linear components of the external finance function,  $\lambda_0$  and  $\lambda_1$ , because the value of financial flexibility increases as external finance becomes more costly. However, the relation shown in the fifth panel between the quadratic component,  $\lambda_2$ , and cash holdings is only slightly upward sloping. With  $\lambda_0$  and  $\lambda_1$  set to their baseline levels, the effect of  $\lambda_2$  is second-order.<sup>4</sup> These results mirror those in the two-period model of Almeida, et al. (2004), which produces a partial derivative of cash with respect to internal funds that is positive for a financially constrained firm, and zero otherwise.

Finally, we examine the effects of technology in the sixth through eighth panels. The sixth panel shows that the effect of  $\theta$  on cash is hump-shaped, initially rising and then falling slightly. Two different economic forces create this pattern. First, as  $\theta$  rises, the production function becomes flatter, and the average size of desired investments rises. The firm holds more cash because large investments imply a greater likelihood of needing external finance. Second, as  $\theta$  rises, the firm is less likely to have to tap external finance because a higher  $\theta$  implies that a given capital stock can create more internal revenue, and the firm therefore needs to hold less cash. The first effect is stronger for lower levels of  $\theta$ , and the second effect is stronger for higher levels of  $\theta$ . The seventh panel shows that cash holding increases with the fixed cost of adjustment. This effect occurs because higher fixed adjustment costs lead to larger investments that occur less frequently. The firm then uses episodes of inaction to accumulate cash, which acts to lower the probability of the firm having to tap external finance when it does invest.<sup>5</sup> Finally, the eighth panel shows the effect of an increase in the convex component of adjustment costs. Not surprisingly, convex adjustment costs have the opposite effect on cash holding. As  $a$  increases, the firm makes smaller investment more often, is therefore less likely to have to tap external finance, and holds less cash.

These results on the level of cash balances reassuringly confirm those in Gamba and Triantis (2006), in particular their results on the effects of uncertainty and the cost of external finance. Our

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<sup>4</sup>This last result may possibly be an artifact of the presence of quadratic physical adjustment in the model because the latter would swamp the former. However, these two sets of costs appear to be operating on different margins because turning off the quadratic physical adjustment costs has almost no effect on the pattern observed for  $\lambda_2$ .

<sup>5</sup>Parenthetically, this type of phenomenon is consistent with the recent tendency in the United States and Europe for firms to hoard cash, as firms concentrate less on internal organic growth than on growth via large acquisitions.

results on cash levels are also useful in providing intuition for the main focus of this paper, which is not cash levels, per se, but the propensity to save.

Figure 2 is analogous to Figure 1, except that it depicts the dependence on the model parameters of our second measure of cash policy—the coefficient  $\alpha_1$  in (10); that is, sensitivity of saving to cash flow. The first panel shows our most interesting simulation result, which is the effect of  $\rho$  on  $\alpha_1$ . For a highly negatively correlated shock process, this sensitivity is large and positive; for a highly positively correlated shock process, this sensitivity is large and negative; and for a predominantly serially uncorrelated income process, this sensitivity is small and negative. This pattern arises out of the firm’s expectation about future needs to tap external finance and about the future productivity of capital. For example, if profits are negatively serially correlated, then if a firm experiences a positive shock, it knows that it will likely experience an income shortfall in the future and consequently be forced to find outside finance for any desired investment projects. It therefore saves. Further, it knows that the productivity of capital goods will decline, and in the face of costly capital reversibility, it will therefore funnel any extra cash flow into saving rather than into investment. On the other hand, if the firm faces positively correlated income shocks, then when it experiences a positive shock, it knows that it will be unlikely to need external finance in the future, and it reduces its cash holdings.<sup>6</sup> Also, because it expects the future productivity of capital to rise, it will funnel extra cash flow into investment rather than saving.

The second panel illustrates the effect of the shock variance. Saving sensitivity is always negative, but becomes less so as  $\sigma_v$  increases. As the firm’s environment becomes more uncertain, its level of cash holdings increases, but it also becomes more reluctant to change its cash holdings aggressively in response to shocks, which convey little information in an uncertain environment. This panel does not imply that the firm always dissaves. Instead, the firm saves and dissaves to counteract profit shocks.

We now turn to the effects of the cost of external finance. The patterns evident in the third through fifth panels mirror those in the corresponding panels in Figure 1. In all cases, as the cost

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<sup>6</sup>In a more fundamental information-theoretic model of the cost of external finance, a positive shock to the firm might induce recontracting between the firm and its supplier of funds. The resulting contract would likely be more favorable to the firm, which would reinforce its tendency to dissave.

of external finance increases, the level of cash holdings increases, and the sensitivity of saving to cash flow becomes more negative. The firm accumulates financial slack when the cost of external finance rises. It can therefore respond to shocks more aggressively by changing the level of cash. For example, if a firm with a great deal of financial slack is hit by a positive profit shock, it will dissave a great deal in order to invest. An otherwise identical firm with little financial slack will not be able to dissave as much.

Finally, we examine technology. Saving sensitivity becomes more negative as  $\theta$  increases; that is, as the production function becomes flatter. With a flat production function shocks induce large desired changes in the capital stock, and the firm dissaves to fund these investments. Saving sensitivity also becomes more negative as both types of adjustment costs increase. As higher adjustment costs cause investment policy to become less flexible, the firm compensates by making its saving policy more flexible.<sup>7</sup>

The preceding arguments are valid at points in time in which the firm is actively adjusting its capital stock. During periods of inaction, the sensitivity of saving to cash flow is positive because the firm funnels at least part of its cash flow into cash holdings in order to avoid tapping external finance in the future. Under almost all parameterizations of this model the firm is adjusting more often than it is remaining inactive. The observations in which the saving sensitivity is negative therefore outweigh those in which it is positive, and average sensitivity is negative. If the firm faces very large fixed adjustment costs, however, it is inactive more often than it invests, the observations with positive sensitivity outweigh those with negative sensitivity, and average sensitivity is positive. We view this latter scenario as mostly likely for very small firms because they are the only ones whose investment tends to occur in large spikes.

The frequent adjustment in our model sets it apart from models of dynamic capital structure with adjustment costs. For example, in Fischer, Heinkel, and Zechner (1989), the firm adjusts its asset and liability composition infrequently. As pointed out in Strebulaev (2006), empirical predictions from this sort of model cannot be based solely on firm behavior at points in time at

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<sup>7</sup>Two parameters we have left out of the analysis are the depreciation rate,  $d$ , and the interest rate,  $r$ . Decreasing the depreciation rate or increasing the discount rate lowers the average size of investments, the need for external finance, and cash levels. The saving propensity remains negative but decreases in absolute value.

which the firm is active. The frequent adjustment in our model allows us to sidestep this critique. Frequent adjustment also explains why corporate propensities to save can be negative even though personal propensities to save are typically positive. Although consumers dissave when they purchase durables, these events are infrequent, and because consumers save out of income at other times, average saving propensities are therefore positive.

Because we are interested in the effect of measurement error in observed Tobin's  $q$  in our data, we conduct a further simulation in which we introduce an additive *i.i.d.* measurement error to  $V(k, p, z)/k$ . Measurement error biases the coefficient on  $V(k, p, z)/k$  downward, but it biases the coefficient on  $\pi(k, z)/k$  upward because of the strong positive correlation between  $\pi(k, z)/k$  and  $V(k, p, z)/k$ . We find that the baseline simulation requires a great deal of measurement error to reverse the initially negative sign of the coefficient on cash flow. The error variance needs to be at least 15 times as large as the variance of  $V(k, p, z)/k$ . This result is not out of line with the empirical results that follow inasmuch as we find that the measurement quality of observed Tobin's  $q$  is extremely low, or equivalently, that the measurement error variance is high.

In sum, these experiments highlight two important pieces of economic intuition. First, corporate saving depends not only on the firm's financial environment, but also on its technological environment. Second, although the levels and changes in cash are clearly related, a high cash level does not necessarily imply a high positive sensitivity of saving to cash flow; nor does a low cash level imply a low sensitivity. We emphasize again that this distinction is impossible to uncover in a static model because the change in cash cannot be distinguished from the level of cash.

### 2.3. Empirical Predictions

The simulations in Figure 2 delineate the four central empirical predictions we wish to test. First, the sign of  $\alpha_1$  in (10) should be negative. Because this prediction takes the form of the sign of a slope coefficient in a linear projection, and because the error term in a linear projection is by definition orthogonal to the right-hand-side variables, testing the prediction with an exactly equivalent real-data linear projection (regression) avoids the usual simultaneity problems that plague regressions in corporate finance. Testing this prediction in this manner also forms a strong link between the theory and its test, because the form of the real-data test is identical to the form of the simulated-

data theoretical prediction. Second,  $\alpha_1$  should increase in absolute value with the cost of external finance; third,  $\alpha_1$  should decrease in absolute value with  $\sigma_v$ ; and fourth,  $\alpha_1$  should increase in absolute value with  $\rho$ . We do not directly test any predictions concerning the relation between  $\alpha_1$  and either profit function curvature or adjustment costs because proxies for these two model features are unavailable at the firm level. The simulations that examine curvature and adjustment costs do, however, provide intuition that assists with the interpretation of some of our results.

## 2.4. Model Robustness

The model is intentionally sparse to highlight the intuition behind the interaction between saving and investment, as well as the trade-off between the tax penalty on saving and the cost of external finance. To assuage concerns that our results are an artifact of the model's simplicity, in this section we add a variety of more realistic features to the model to examine the robustness of our result of a negative propensity to save. First, in the baseline model the firm does not have access to a credit line. When we add riskless short-term debt that is secured by the capital stock, as in Hennessy and Whited (2005), the saving propensity of -0.78 in the baseline simulation drops in absolute value to -0.41. Our results are attenuated but not erased because the upper limit to the credit line causes cash to retain its value as a tool to avoid costly external finance. Second, the baseline firm does not smooth distributions to shareholders. To address this possibility, we add a linear cost if the firm deviates from a target distribution. The cost function takes the form  $0.05 |e - 0.13k| (1 - \Phi_e)$ , in which 0.13 is the average ratio of distributions to the capital stock in the baseline simulation, and the coefficient 0.05 is set to a conservative value. This model feature produces increased cash hoarding because the firm wants to avoid missing a distribution. This higher cash cushion leads to a more negative propensity to save of -0.86. Third, the baseline firm has no fixed costs of production, which could, for example, represent the tendency of young firms to burn profits. We add a cost of production equal to  $0.4k$ , in which 0.4 is the approximate ratio of selling, general, and administrative expenses to assets in our U.S. sample. This addition to the model produces less cash holding relative to our baseline model because the firm has smaller profits to funnel into liquid assets. Accordingly, the saving propensity drops in absolute value to -0.26. Fourth, any real-world cross section of firms contains firms with heterogeneous technologies. To

simulate this situation, we divide our panel of simulated firms into ten groups, each of which has a value of  $\theta$  with one of ten equally spaced values between 0.6 and 0.9. For this simulation we find that the saving propensity is slightly lower in absolute value at -0.53. Sixth, we consider the possibility that the firm may issue risky debt, which we model exactly as in Hennessy and Whited (2006). This approach necessitates the exclusion of any physical adjustment costs. In this case the firm hoards more cash than in the baseline simulation in order to avoid default, and the saving propensity rises in absolute value to -0.88. Finally, because our model contains only one source of uncertainty, productivity shocks and cash flow are almost perfectly correlated. Therefore, the dissaving that occurs with a positive productivity shock is necessarily accompanied by a rise in cash flow. To ascertain whether our finding of a negative saving propensity is hardwired by this feature of our model, we allow the revenue function to take the form  $zk^\theta - \eta$ , in which  $\eta$  is a normally distributed, zero-mean, *i.i.d.* random variable with a variance equal to that of the  $z$  shock. This new revenue shock takes four points of support, and its transition matrix is given by the method in Tauchen (1986). Not surprisingly, decoupling cash flows from the productivity shock produces a much smaller (in absolute value) propensity to save. It is, however, still negative at a value of -0.15. We choose to exclude these features from our baseline model because they do not change the qualitative outcomes of the simulation and because the empirical emphasis of the paper lends itself to a simple structure that stresses intuition.

### 3. Data and summary statistics

We draw data for nonfinancial firms from Standard and Poor's Compustat Global Issue and Industrial/Commercial for six countries: Canada, France, Germany, Japan, the United Kingdom, and the United States. Data cover the period 1997-2004. We select the sample by first deleting any firm-year observations with missing data. Next, we delete any observations for which total assets, the gross capital stock, or sales are either zero or negative. Then for each firm we select the longest consecutive times series of data and exclude firms with only one observation. Finally, we omit all firms whose primary SIC classification is between 4900 and 4999, between 6000 and 6999, or greater

than 9000, because our model is inappropriate for regulated, financial, or quasi-public firms.<sup>8</sup>

We define data variables from Global Industrial/Commercial as follows: Book assets is Item 89; investment is Item 193; operating income is Item 14; cash flow is the sum of Items 11 and 32; and cash is Item 60. The numerator of the market-to-book ratio is the sum of the market value of equity (Item 3 times Item 13 in Global Issue) and total book assets minus the book value of equity (Item 105+Item 135), and the denominator is book assets. In our regressions we scale both saving and cash flow by total assets. We delete the top and bottom 1% of our regression variables.

The number of observations available varies across years, and these numbers are shown in Tables 2-7 by year and country. Summary statistics are in Table 1. We see large differences in most instances between our means and medians. This skewness is essential for identifying our econometric model. The mean and median measures of Tobin's  $q$  (market to book) are greatest in the United States and lowest in Japan and France. Not surprisingly, the investment rates in Japan and France are low. Only France shows negative mean investment, but its median investment level is positive, though small compared with investment in the other countries. All means and medians of Tobin's  $q$  are greater than 1. Although this pattern is commonly viewed as an indication of positive investment opportunities, adjustment and installation costs in the investment process and market power can leave  $q$  above 1 even for firms without positive investment opportunities. The average change in the cash stock is negative in the United Kingdom, Japan, France and Germany, whereas it is positive in the United States and Canada.

## 4. Estimation

This section first outlines the econometric methodology; details are provided in the Appendix. It next presents the results and their interpretation.

### 4.1. Methods

Our testing strategy is based on the estimators in Erickson and Whited (2000, 2002). These estimators employ the structure of the classical errors-in-variables model. Applied to a single cross

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<sup>8</sup>Estimation of our model with data from Thompson Financial's Worldscope produces results similar to those reported here. We omit these because missing data in the Worldscope reduces sample size. We also choose not to blend the two data sets to preserve measurement consistency in our reported results.

section, this model can be written as

$$y_i = w_i\alpha + \chi_i\beta + u_i, \tag{11}$$

$$x_i = \gamma + \chi_i + \varepsilon_i. \tag{12}$$

In our application  $y_i$  is the ratio of the change in cash to assets,  $\chi_i$  is the true  $q$  of firm  $i$ ,  $x_i$  is an estimate of its true  $q$ , and  $w_i$  is a row vector of perfectly measured regressors, whose first entry is 1 and whose second entry is the ratio of cash flow to assets.<sup>9</sup> The regression error,  $u_i$ , and the measurement error,  $\varepsilon_i$ , are assumed to be independent of each other and of  $(w_i, \chi_i)$ , and the observations,  $(\varepsilon_i, u_i, w_i, \chi_i)$ ,  $i = 1, \dots, n$ , are *i.i.d.* The intercept,  $\gamma$ , in (12) allows for systematic bias in the measurement of  $\chi_i$ . We do not require any assumptions about the temporal dependence or independence of  $(\chi_i, w_i, u_i, \varepsilon_i)$ .

Using the third and higher order moments of  $(x_i, y_i)$ , the Erickson and Whited estimators provide consistent estimates of the slope coefficients,  $\alpha$  and  $\beta$ , as well as the variances of the unobservable variables  $(\chi_i, u_i, \varepsilon_i)$ . These estimators are identified only if  $\beta \neq 0$  and  $\chi_i$  is nonnormally distributed. Erickson and Whited (2002) develop a test of the null hypothesis that  $\beta = 0$  and  $\chi_i$  is nonnormally distributed—a test we refer to hereafter as an identification test. For details see Erickson and Whited (2000, 2002).

The Appendix presents Monte Carlo simulations to assess the finite sample performance of these estimators on data closely resembling our own. Of particular interest in these Monte Carlos are the tests of the null hypothesis that the coefficient on cash flow equals its true value. The actual sizes of many of these tests are tiny relative to their nominal sizes. This result indicates that the finite-sample distribution of the GMM estimates of the t-statistics has thin tails. This result on test-size is the opposite of that found in Erickson and Whited (2000) for *investment* regressions. The large discrepancy between the actual and nominal test sizes prompts us to use bootstrapped critical values for the t-statistics on these coefficients, although using asymptotic critical values leads us to most of the same conclusions. We use the GMM bootstrapping procedure in Brown and Newey (2002), which resamples from the empirical likelihood distribution that imposes the moment

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<sup>9</sup>Almeida et al. (2004) also include the log of firm size in (10). We omit size because it is never significant in our empirical work and because it has no impact on our results.

restrictions.

Because these estimators can only be applied to samples that are arguably *i.i.d.*, we estimate (11) and (12) for each cross section of our unbalanced panel. We do not include firm fixed effects in our regressions for two reasons. First, the resulting model almost never passes the identification test. Second, our OLS results are little changed by the inclusion of fixed effects, suggesting that the within-firm variation in investment and  $q$  mirrors the cross sectional variation.<sup>10</sup> As an informal summary of our yearly results, we pool the yearly estimates from our unbalanced panel via the procedure in Fama and MacBeth (1973). Recently, Petersen (2005) has re-emphasized that Fama-Macbeth standard errors are often inappropriate in panel data. However, we report these standard errors only as an informal summary of our results and instead focus on the GMM standard errors for each cross section. Interestingly, the two sorts of standard errors lead to the same inference.

## 4.2. Results

Before discussing the results, we emphasize that we fail to reject the overidentifying restrictions from our yearly GMM estimates in all countries for most years. This result mitigates concerns about possible model misspecification. For example, although  $u_i$  is by construction uncorrelated with  $(\chi_i, w_i)$ , it may not be independent of  $(\chi_i, w_i)$ . Similarly,  $\varepsilon_i$  may not be independent of  $(\chi_i, w_i)$ , or the true form of the regression (11) may be nonlinear. Nonetheless, even though the classical errors-in-variables model is not a perfect representation of the relationship between saving, cash flow, and Tobin's  $q$ , our tests indicate that we have a useful approximation. Our failure to reject is even more interesting given the result in the Appendix that this test has a slight tendency to over-reject in finite samples.

Results from estimating (11) via OLS and from estimating (11) and (12) via GMM are in Tables 2 through 7, by country. In each table the left panel shows the OLS results, and the right panel shows the GMM results. We report the OLS estimate of the coefficient of determination of (11), (the regression  $R^2$ ) in column 3, and the measurement-error consistent GMM estimate of the coefficient of determination of (11) (denoted  $r^2$ ) in column 6. Column 7 contains the coefficient

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<sup>10</sup>Because there may nonetheless be some cross-sectional dependence among the firms in the same industry, we also try to include two-digit industry fixed effects. The results are quite similar, though we find a higher incidence of unidentified models and consequently higher standard errors.

of determination of (12) (denoted  $\tau^2$ ), which is a useful index of measurement quality for our observable proxy for unobservable investment opportunities. A value close to 1 indicates a nearly perfect proxy, and a value close to zero indicates a nearly worthless proxy. The eighth column reports the  $J$ -test of the overidentifying restrictions of the model. The null hypothesis of this test is that the overidentifying restrictions are satisfied. The final column contains the identification test results. The null in this case is that the model is unidentified. A number of researchers have used this test as a “pre-sampling” test to select samples on which the estimator is identified. We do not follow this practice. Rather, we select our sample based on the minimal criterion of data availability. A failure to reject the null of the identification test simply means that the estimates are unreliable—a tendency reflected in large standard errors.

Asymptotic standard errors are in parentheses below each parameter estimate, and p-values are in parentheses below the chi-squared tests. Asterisks mark those GMM estimates of the cash-flow coefficient whose t-statistics exceed the 5% bootstrapped critical values. For brevity, we only indicate the significance of the cash-flow coefficients.

For each country we test our first prediction that the coefficient on cash flow is negative in a regression of saving on  $q$  and cash flow. Our OLS results corroborate earlier findings that the coefficients on both Tobin’s  $q$  and cash flow are positive in virtually all years for all countries, as in, e.g., Almeida, et al. (2004) and Khurana, et al. (2006). However, when we apply the Erickson and Whited estimators to correct for measurement error in  $q$ , the results change. We find negative coefficients on cash flow in virtually all cases in all six countries, and over half of the cash-flow coefficients are statistically significant according to our bootstrapped critical values. These results correspond to our model simulation results for firms that have positively serially correlated income processes. Sample size almost assuredly affects our ability to find statistical significance when using the GMM estimator inasmuch as a small sample size renders the GMM model unidentified and can, therefore, weaken results. This problem occurs even though Tobin’s  $q$  is a price and therefore, by definition, has a skewed distribution because detecting this skewness requires a great deal of data. As is common when using international data, in some countries—particularly Canada, France, and Germany—we have smaller samples than we would like. This problem is especially evident in Table

6, which contains the results for France. Given these difficulties, our finding of a high incidence of significant coefficients is all the more striking.

The effect of treating measurement error can also be seen in the GMM coefficients on  $q$ , which are from seven to fifteen times as high as their OLS counterparts in the different countries. This result can be explained by the attenuation bias in the classical errors-in-variables model, combined with our low point estimates for  $\tau^2$ , which hover around 0.1. The attenuation bias is given by  $\text{plim}(\hat{\beta}_{OLS}) = \beta\tau_x^2$ , in which  $\tau_x^2 \equiv [(\tau^2 - R_x^2)/(1 - R_x^2)]$ , with  $R_x^2$  being the probability limit of the sample coefficient of determination from the OLS regression of  $x_i$  on  $w_i$ .

Next we test our second prediction that the cash-flow coefficient is more negative for firms with more costly external finance. To this end we split our samples according to firm size. We define a firm as large if the book value of its assets lies above the sample median in the country in question; otherwise, we define a firm as small. In the literature on external finance constraints, size is often used as an indicator of the cost of raising external funds. Using size as an indicator of the cost of external finance confers an important advantage over other indicators such as dividend payout. It can be considered exogenous, because firm size is not a choice variable for the manager in the short run and because our estimates exploit cross sectional data variation instead of time series variation.

Tables 8 and 9 report results for the United States and Japan. Sample sizes in the remaining countries are too small to allow GMM estimation with a split sample. We begin with the results for the U.S. split sample. Our OLS results confirm those in Almeida, et al. (2004) that small U.S. firms' saving is more positively related to cash flow than is large U.S. firms' saving. Our GMM estimates, however, are still predominately negative for both the large and small firms. Nevertheless, these two subsamples do exhibit important differences. We find a mix of positive and negative coefficients for the small firm subsample, and the average coefficient is -0.033. In contrast, for the large firms we find that many more of the GMM estimates of the yearly cash-flow coefficients are negative and significant, and the average is quite large at -0.692.<sup>11</sup>

Our results for the United States do not support the model's prediction that firms with more

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<sup>11</sup>Because both saving and investment decisions respond less to shocks as uncertainty increases, this result is consistent with those in Kadapakkam, Kumar and Riddick (1998), who find that investment-cash flow sensitivity is lower in large firms than in small firms.

costly external finance have more negative cash-flow coefficients. Because of the evidence in Hennessy and Whited (2006) that small firms do indeed face more costly external finance than do large firms, we conjecture that the effect of the cost of external finance on saving operates in our sample but that the effect of uncertainty is stronger. When we examine differences between the large and small firms, one characteristic that stands out is the marked difference in the degree of uncertainty that they face. We estimate a first-order panel autoregression of the ratio of operating income to assets for both groups of firms, using the technique in Holtz-Eakin, Newey, and Rosen (1988). The standard deviation of the error term for the small firms is 0.140, whereas for the large firms it is only one third as large at 0.046. As demonstrated in the model simulations, firms that face a great deal of uncertainty do not make large changes in their cash holdings in response to income shocks.

Our conjecture that uncertainty outweighs the cost of external finance is also supported by a model simulation in which we triple  $\sigma_v$  and double  $\lambda_0$ ,  $\lambda_1$ , and  $\lambda_2$  relative to the baseline case. Tripling  $\sigma_v$  is consistent with the difference in uncertainty between large and small firms in our U.S. sample; and doubling the external finance parameters is consistent with the results in Hennessy and Whited (2006), who find that external equity costs for small firms are about twice those for large firms. We find that the cash-flow coefficient drops after this change because the effect of uncertainty is dominant.

It is possible that the low sensitivity in the small firms may be due to the tendency of the investment of very small firms to occur infrequently, but in large spikes. During periods of inactivity the firm saves out of cash flow to avoid costly external finance when it does invest. However, we view this possibility as unlikely. In only 5.7% of the firm-year observations in this sample do we observe investment rates greater than twice the firm median. For further discussion of investment spikes in firm-level U.S. data, see Whited (2006).

The results for Japan are less pronounced. Although we do find that most of the cash-flow coefficients are more negative for large firms than for small firms, the difference between the coefficients in the two subsamples is smaller. In addition, the difference between the levels of uncertainty that these two groups of Japanese firms face is markedly smaller than in the case of the United States. The standard deviations of the error in a first-order autoregression of operating income for

the small and large firms are 0.042 and 0.035, respectively. We conjecture that these low levels of uncertainty may occur because of risk-sharing among firms in the keiretsu structure.

It is natural at this point to test our third and fourth predictions; that is, to investigate directly the relation between saving propensities and the serial correlation and variance of operating income. To this end we estimate an  $AR(1)$  for operating income (scaled by total assets) firm by firm, only using firms with at least six consecutive observations. We then sort our samples from the United States and Japan on the basis of the estimates of serial correlation and residual standard deviation, throw out the middle third, and compare saving propensities across the top and bottom thirds. We discard the middle third because firm-by-firm estimates contain a great deal of noise, and we want to minimize the possibility that we incorrectly classify individual firms.<sup>12</sup>

The results from the regressions that examine low and high serial correlation for U.S. firms are in Table 10. Interestingly, both the OLS and GMM estimates of the coefficient on cash flow are positive for the low serial correlation firms, although the yearly GMM estimates are smaller than the yearly OLS estimates and, for the most part, insignificantly different from zero. In this subsample the median serial correlation is near zero. In contrast, for the high serial correlation firms the GMM estimate of the cash-flow coefficient is large and significantly negative. In this subsample the median serial correlation is 0.72. These results support the simulation that shows a low and sometimes positive saving sensitivity if the serial correlation of the income process is low, but a large negative saving sensitivity if the serial correlation is high. The results for Japan in Table 11 are similar, except that the GMM estimates of the cash-flow coefficient for the high serial correlation group are, although large and negative, not significantly different from zero. The median serial correlations for the two groups are -0.035 and 0.677. None of these results are driven by differences in size across groups. The mean of total assets is nearly identical across groups in both countries.

Tables 12 and 13 contain parallel results for firms grouped by the standard deviation of the residual of the  $AR(1)$  process for operating income. As predicted by the model, the low standard

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<sup>12</sup>Noise in the firm-level estimates of serial correlation and residual standard deviation means that we cannot include these estimates in our regression. In this case the regression would contain more than one mismeasured regressor, and we do not have enough data to use a two-mismeasured-regressor version of the Erickson and Whited estimators.

deviation group has a significantly negative cash-flow coefficient, whereas the high standard deviation group has a near-zero, insignificant cash-flow coefficient. Not surprisingly, given the results on large and small U.S. firms, the low standard deviation group contains firms considerably larger than those in the high standard deviation group. The mean level of assets for the former is 3.939 million 1997 dollars, and the mean level of assets for the latter is 1.074 million 1997 dollars. The results for Japan are indeterminate. Both groups have negative cash-flow coefficients of similar magnitude. We attribute this result to the low dispersion of residual standard deviations for firms in Japan. When we divide the sample, we do not find enough variation in the residual standard deviation to obtain groups of firms that differ much from one another. Indeed, the average level of total assets in the two groups is almost identical.

As an informal test, we next investigate the serial correlation and residual standard deviation of operating income across countries. For this exercise we again use the panel-VAR approach in Holtz-Eakin, Newey, and Rosen (1988) to estimate a first-order autoregressive model, finding statistically significant autoregressive coefficients of 0.55, 0.43, 0.57, 0.46, 0.46, and 0.65 for the United States, Canada, the United Kingdom, Japan, France, and Germany, respectively. This evidence of positive serial correlation is reassuring, given our simulation results that positive serial correlation is associated with a negative cash-flow coefficient. The cross-sectional variation in these autoregressive coefficients is, however, sufficiently small that the association between them and the average cash-coefficients is weak, although, as predicted, negative. Such is not the case with the estimates of the error standard deviations from these auto regressions. These estimates are 0.115, 0.107, 0.092, 0.038, 0.070, and 0.068 for the United States, Canada, the United Kingdom, Japan, France, and Germany, respectively. In this case we find a positive correlation of 0.64 between the error standard deviations and the average cash-flow coefficients. Although it is difficult to draw strong inferences from six data points, this result is all the more striking given that the correlation coefficient is statistically significant at the 10% level. The result confirms the intuition from the model that if a firm is subject to a noisy income process, it tends to change its cash holdings less aggressively in response to movements in income.

One wonders if the degree of uncertainty that is so important in understanding these results

is merely proxying for financial constraints. It is possible that external finance is more costly for firms that face more uncertainty. However, two pieces of evidence indicate that this possibility does not drive our results. First, if uncertainty were proxying for the costliness of external finance, we ought to have seen a more negative cash-flow coefficient for small firms than for large firms. We also ought to have seen a more negative coefficient for firms with high residual standard deviations than for firms with low residual standard deviations. We do not. Second, we find a relationship across countries between uncertainty and saving sensitivity, and it is unlikely that this result is due to differential access to external finance in the six highly developed countries we examine.

To conclude, we compare our results with those in Erickson and Whited (2000), who find that the measurement-error consistent estimates of the coefficients on cash flow in a regression of investment on  $q$  and cash flow are near zero. As explained in the introduction, measurement error in a variable biases other coefficients via the correlation between the mismeasured variable and the other regressors. Not required for bias is a correlation between the measurement error and other regressors. The regressions in this paper and those in Erickson and Whited (2000) have the same sets of right-hand-side variables and only differ in the dependent variable. It seems odd, therefore, that our results are different. However, the measurement quality of the proxy for investment opportunities used in the saving literature, the market-to-book ratio, is of particularly low quality; most of the estimates of the index of measurement quality,  $\tau^2$ , cluster around 0.1. In contrast, the measurement quality of the proxy used by Erickson and Whited (2000) is approximately 0.4. Therefore, the positive OLS estimates of the cash-flow coefficients in our regressions are biased upward much more severely than the positive estimates in Erickson and Whited (2000). Correcting for measurement error in both papers produces consistent estimates of the cash-flow coefficients, which are negative here and near zero in Erickson and Whited (2000).

To elaborate we turn to an identity that expresses the cash-flow coefficient as a function of the coefficient on  $q$ . The coefficient  $\alpha_1$  on cash flow in (10) can be expressed as

$$\alpha_1 = \mu_{1y} - \beta\mu_{1x},$$

in which  $(\mu_{1y}, \mu_{1x})$  are the slope coefficients obtained from projecting saving and observed  $q$  on cash flow, and  $\beta$  is the coefficient on  $q$ . On average in our regressions  $\mu_{1y}$  is about 0.15, and

$\mu_{1x}$  is about 1.5. The biased OLS estimate of  $\beta$ , which hovers around 0.02, therefore produces a positive cash-flow coefficient when plugged into the above identity with these values for  $\mu_{1y}$  and  $\mu_{1x}$ . In contrast, the consistent GMM estimate of  $\beta$ , which hovers around 0.2, produces a negative cash-flow coefficient.

In Erickson and Whited (2000) the values of  $\mu_{1y}$  and  $\mu_{1x}$  are higher than in our application, and the bias in the OLS estimates of  $\beta$  is not as severe because of the higher measurement quality of the  $q$  proxy. These numbers, when plugged into the above identity therefore produce positive OLS coefficients on cash flow and near-zero GMM coefficients. Not surprisingly, in results not reported for brevity, we are unable to find positive OLS cash-flow coefficients in our saving regressions when using the proxy from Erickson and Whited (2000); and, as is the case here, we find negative GMM estimates of the cash-flow coefficients.

## 5. Conclusion

The issue of corporate saving has recently received much attention, in large part because of the tendency in recent years of both U.S. and European firms to accumulate a great deal of liquid assets. Prior empirical research, including papers by Opler, Pinkowitz, Stulz, and Williamson (1999), Almeida, et al. (2004), Faulkender and Wang (2006), and Khurana, et al. (2006), has addressed two related issues: why firms hold cash and why firms save, that is, change their cash holdings. We have, for the most part, addressed the second issue. In so doing, we take care to model a firm's saving, financing, and real investment decisions simultaneously in a stochastic, dynamic framework, to form a strong link between our theory and our empirical tests, and to account for measurement error in our empirical work.

This approach leads us to conclusions quite different from previous theoretical and empirical results. Our theoretical model predicts that in the face of costly external finance, uncertainty in income, and taxation, the firm counteracts movements in cash flow with opposite movements in saving. This negative propensity to save occurs because the firm wants to funnel excess cash flow into capital goods rather than into cash, after it receives good cash-flow news. In contrast, the sensitivity of saving to cash flow in earlier static models, such as the one in Almeida, et al. (2004),

is typically positive because the firm wants to hold high levels of cash to avoid costly external finance. This latter effect operates in our model as well. However, in our model this effect pertains to the level of cash rather than to the change in cash (saving). These distinctions between the level of cash and saving are indistinguishable in a static model, and earlier work could not, therefore, reach our results.

We find strong empirical support for our negative-sensitivity result in data from six countries. When we initially examine this sensitivity using OLS regressions, we find the standard result in the literature that the sensitivity of saving to cash flow is positive. However, when we correct econometrically for measurement error in Tobin's  $q$ , we find the opposite result. We also find that the sensitivity of saving to cash flow increases in absolute value with the serial correlation of income and decreases with the variance of income shocks. Interestingly, both of these effects on saving are empirically stronger than the effect of financial constraints. A natural consequence of this result is that propensities to save cannot be used as summary measures of the cost of external finance.

Taken together, our simulations and empirical evidence demonstrate that income shocks are at least as important as financial constraints in determining corporate saving. Our results also reemphasize that researchers should determine if any model estimation involving Tobin's  $q$  has similarly significant bias from mismeasurement. The variable  $q$  is important in many contexts, and its mismeasurement has the potential to bias results in any context where  $q$  is correlated with other variables in the model. Because  $q$  is a broad measure of firm health, it is likely that this correlation issue will be important in many situations, as it is here.

## 6. Appendix. Estimation and Monte Carlo Experiments

To derive a set of tractable moment conditions that can be estimated with GMM, the vector of perfectly measured variables,  $w_i$ , must be partialled out of equations (11) and (12). Let  $(\dot{y}_i, \dot{x}_i, \dot{\chi}_i)$  be the residuals from the linear projection of  $(y_i, x_i, \chi_i)$  on  $w_i$ . Then (11) and (12) can be written as

$$\dot{y}_i = \beta \dot{\chi}_i + u_i \quad (13)$$

$$\dot{x}_i = \dot{\chi}_i + \varepsilon_i. \quad (14)$$

If we square (13), multiply the result by (14), and take unconditional expectations of both sides, we obtain

$$E(\dot{y}_i^2 \dot{x}_i) = \beta^2 E(\dot{\chi}_i^3). \quad (15)$$

Analogously, if we square (14), multiply the result by (13), and take unconditional expectations of both sides, we obtain

$$E(\dot{y}_i \dot{x}_i^2) = \beta E(\dot{\chi}_i^3). \quad (16)$$

As shown in Geary (1942), if  $\beta \neq 0$  and  $E(\dot{\chi}_i^3) \neq 0$ , dividing (15) by (16) produces a consistent estimator for  $\beta \equiv \beta^2 E(\dot{\chi}_i^3) / \beta E(\dot{\chi}_i^3)$ . The assumptions,  $\beta \neq 0$  and  $E(\dot{\chi}_i^3) \neq 0$ , are necessary for identification. They can be tested via the null hypothesis that  $E(\dot{y}_i^2 \dot{x}_i) = 0$  and  $E(\dot{y}_i \dot{x}_i^2) = 0$ .

This estimator is a third-order moment estimator. The innovation in Erickson and Whited (2000, 2002) consists of combining the information in moment equations of orders two through seven via GMM to obtain a more efficient estimator for  $\beta$ . The coefficient  $\alpha$  in can be recovered by the identity

$$\alpha = \mu_y - \beta \mu_x,$$

in which  $(\mu_y, \mu_x)$  are the vectors of coefficients in the population projection of  $(y_i, x_i)$  on  $w_i$ . The coefficients of determination ( $R^2$ 's) for (11) and (12) are calculated as

$$r^2 = \frac{\mu_y' \text{var}(w_i) \mu_y + E(\dot{\chi}_i^2) \beta^2}{\mu_y' \text{var}(w_i) \mu_y + E(\dot{\chi}_i^2) \beta^2 + E(u_i^2)} \quad (17)$$

$$\tau^2 = \frac{\mu_x' \text{var}(w_i) \mu_x + E(\dot{\chi}_i^2)}{\mu_x' \text{var}(w_i) \mu_x + E(\dot{\chi}_i^2) + E(\varepsilon_i^2)}. \quad (18)$$

In order to allay skepticism of empirical results that have been produced by unusual estimators on fairly small samples, in Table 14 we report the results of a Monte Carlo simulation using artificial data similar to our real data, both in terms of sample size and observable moments. These simulations are of particular interest because these estimators have most commonly been used on investment regressions instead of saving regressions, and because saving and investment have different statistical properties. Most importantly, the distribution of investment is highly skewed, whereas the distribution of saving is much more symmetric.

We do three experiments. For each we generate 10,000 simulated cross sections from (13) and (14). The first has a sample size of 2,576, the second a sample size of 1,200, and the third a sample size of 208. These numbers correspond to the size of the largest and smallest cross sections in our data set, as well as to an intermediate size. For each simulation we set the parameters  $\beta$ ,  $\alpha$ ,  $r^2$ , and  $\tau^2$  approximately equal to the averages of the corresponding GMM estimates from Tables 2 through 7. Each observation is of the form  $(y_i, x_i, w_i)$ , generated according to (11)-(12) so that  $(y_i, x_i, w_i)$  has, on average over the simulation samples, first and second moments equal to, and higher-order moments comparable to, the corresponding average sample moments from our real data.

For the third-, fourth-, fifth-, and sixth-order GMM estimators, Table 14 reports the mean value of the estimator of our parameter of interest,  $\alpha_1$ . It also reports its mean absolute deviation (MAD), the probability that an estimate is within 20% of its true value, and the actual size of a nominal 5% two-sided test of the null hypothesis that  $\alpha_1$  equals its true value. For the small and intermediate sample sizes Table 14 shows that the fourth-order GMM estimator (GMM4) gives the best estimates in terms of expected value, MAD, and probability concentration. For the large sample size the GMM6 estimator performs best. Because the performance of the GMM4 and GMM6 estimators is similar for the large sample size, we therefore use the GMM4 estimator for our empirical work. Also of interest in this table are the tiny actual sizes of the test of the null hypothesis that  $\alpha_1$  equals its true value for the intermediate and large sample sizes.

## References

- Acharya Viral, Heitor Almeida, and Murillo Campello, 2004, Is cash negative debt? A hedging perspective on corporate financial policies, Working paper, New York University.
- Almeida, Heitor, Murillo Campello, and Michael S. Weisbach, 2004, The cash flow sensitivity of cash, *Journal of Finance* 59, 1777–1804.
- Altinkilic, Oya and Robert S. Hansen, 2000, Are there economies of scale in underwriting fees? Evidence of rising external financing costs, *Review of Financial Studies* 13, 191–218.
- Brown, Bryan W. and Whitney K. Newey, 2002, Generalized method of moments, efficient bootstrapping, and improved inference, *Journal of Business and Economic Statistics* 20, 507–517.
- Caggese, Andrea, 2006, Testing financing constraints on firm investment using variable capital, *Journal of Financial Economics*, forthcoming.
- Cooley, Thomas F. and Vincenzo Quadrini, 2001, Financial markets and firm dynamics, *American Economic Review* 91, 1286–1310.
- Cooper, Russell and Joao Ejarque, 2003, Financial frictions and investment: requiem in  $q$ , *Review of Economic Dynamics* 6, 710–728.
- Cooper, Russell and John Haltiwanger, 2006, On the nature of capital adjustment costs, *Review of Economic Studies* 73, 611 - 633.
- Dittmar, Amy and Jan Mahrt-Smith, 2005, Corporate governance and the value of cash holdings, *Journal of Financial Economics*, forthcoming.
- Eisfeldt, Andrea L. and Adriano A. Rampini, 2006, Financing shortfalls and the value of aggregate liquidity, Working paper, Northwestern University.
- Erickson, Timothy and Toni M. Whited, 2000, Measurement error and the relationship between investment and  $q$ , *Journal of Political Economy* 108, 1027–57.
- Erickson, Timothy and Toni M. Whited, 2002, Two-step GMM estimation of the errors-in-variables model using high-order moments, *Econometric Theory* 18, 776–799.
- Fama, Eugene and James MacBeth, 1973, Risk, return, and equilibrium: empirical tests, *Journal of Political Economy* 81, 607–36.
- Faulkender, Michael and Rong Wang, 2006, Corporate financial policy and cash holdings, *Journal of Finance* 61, 1957–1990.
- Fischer, Edwin O., Robert Heinkel, and Josef Zechner, 1989, Dynamic capital structure choice: theory and tests, *Journal of Finance* 44, 19–40.
- Foley, C. Fritz, Jay C. Hartzell, Sheridan Titman, and Garry Twite, 2006, Why do firms hold so much cash? A tax-based explanation, *Journal of Financial Economics*, forthcoming.
- Gamba, Andrea and Alexander Triantis, 2006, The value of financial flexibility, Working paper, University of Maryland.
- Geary, R. C., 1942, Inherent relations between random variables, *Proceedings of the Royal Irish Academy A* 47, 63–76.
- Gomes, Joao, 2001, Financing investment, *American Economic Review* 91, 1263–1285.
- Hennessy, Christopher A. and Toni M. Whited, 2005, Debt dynamics. *Journal of Finance* 60, 1129–1165.

- Hennessy, Christopher A. and Toni M. Whited, 2006, How costly is external financing? Evidence from a structural estimation, *Journal of Finance*, forthcoming.
- Holtz-Eakin, Douglas, Whitney K. Newey, and Harvey Rosen, 1988, Estimating vector autoregressions with panel data, *Econometrica* 56, 1371-1395.
- Kadapakkam, Palani Rajan, P. C. Kumar, and Leigh A. Riddick, 1998, The impact of cash flows and firm size on investment: the international evidence, *Journal of Banking and Finance* 22, 293-320.
- Khurana, Inder K., Xiumin Martin, and Raynolde Pereira, 2006, Financial development and the cash flow sensitivity of cash, *Journal of Financial and Quantitative Analysis*, forthcoming.
- Kim, Chang-Soo, David C. Mauer, and Ann E. Sherman, 1998, The determinants of corporate liquidity: Theory and evidence, *Journal of Financial and Quantitative Analysis* 33, 335-359.
- Moyen, Nathalie, 2004, Investment-cash flow sensitivities: Constrained versus unconstrained firms, *Journal of Finance* 59, 2061-2092.
- Opler, Timothy, Lee Pinkowitz, René Stulz, and Rohan Williamson, 1999, The determinants and implications of corporate cash holdings, *Journal of Financial Economics* 52, 3-46.
- Petersen, Mitchell A., 2005, Estimating standard errors in finance panel data sets: comparing approaches, Working paper, Northwestern University.
- Pinkowitz, Lee, and Rohan Williamson, 2001, Bank power and cash holdings: Evidence from Japan, *Review of Financial Studies* 14, 1059-1082.
- Rotemberg, Julio J. and Michael Woodford, 1992, Oligopolistic pricing and the effects of aggregate demand on economic activity, *Journal of Political Economy* 100, 1153-1207.
- Rotemberg, Julio J. and Michael Woodford, 1999, The cyclical behavior of prices and costs. In: Taylor, J. B., Woodford, M. (Eds.), *Handbook of Macroeconomics*, Vol. 1B. North Holland, Amsterdam, pp. 1051-1135.
- Stokey, Nancy L. and Robert E. Lucas, 1989, *Recursive Methods in Economic Dynamics*. Harvard University Press, Cambridge, Mass. and London.
- Strebulaev, Ilya, 2006, Do tests of capital structure mean what they say? *Journal of Finance*, forthcoming.
- Sufi, Amir, 2006, The real effects of debt certification: evidence from the introduction of bank loan ratings, Working paper, University of Chicago.
- Tauchen, George, 1986, Finite state Markov-chain approximations to univariate and vector autoregressions, *Economics Letters* 20, 177-181.
- Whited, Toni M., 2006, External finance constraints and the intertemporal pattern of intermittent investment, *Journal of Financial Economics* 81, 467-502.

Table 1: Summary Statistics

		Investment	$q$	Cash Flow	Cash Stock	Change in Cash Stock
United States						
	Mean	0.0733	1.7398	0.0655	0.1383	0.0093
	Median	0.0483	1.3869	0.0863	0.0541	0.0013
Canada						
	Mean	0.1202	1.3935	0.0729	0.1026	0.0047
	Median	0.0690	1.1634	0.0898	0.0349	0.0000
United Kingdom						
	Mean	0.0680	1.4619	0.0872	0.1113	-0.0014
	Median	0.0427	1.2630	0.0954	0.0662	0.0000
Japan						
	Mean	0.0269	1.0796	0.0344	0.1620	-0.0043
	Median	0.0183	1.0145	0.0338	0.1344	-0.0026
France						
	Mean	-0.0027	1.2150	0.0613	0.1291	-0.0156
	Median	0.0016	1.0725	0.0644	0.0911	-0.0031
Germany						
	Mean	0.0348	1.2108	0.0673	0.1100	-0.0174
	Median	0.0308	1.1070	0.0729	0.0572	-0.0041

Calculations are based on a sample of nonfinancial firms from Global Vantage from 1997 to 2004.  $q$  stands for the market-to-book ratio. Investment, cash flow, the cash stock, and the change in the cash stock are all deflated by total assets.

Table 2: Cash Regressions: USA

Year	OLS			GMM						
	$q$	$CF$	$R^2$	$q$	$CF$	$r^2$	$\tau^2$	$J$ -Test	ID Test	Obs.
1997	0.039 (0.007)	0.060 (0.036)	0.108 (0.026)	0.263 (0.017)	-0.219* (0.071)	0.615 (0.043)	0.148 (0.033)	3.743 (0.154)	10.139 (0.006)	2576
1998	0.031 (0.007)	0.155 (0.037)	0.107 (0.028)	0.233 (0.024)	-0.120* (0.078)	0.534 (0.075)	0.134 (0.038)	6.185 (0.045)	10.661 (0.005)	2615
1999	0.013 (0.004)	0.197 (0.030)	0.090 (0.019)	0.238 (0.035)	-0.314* (0.130)	0.383 (0.070)	0.056 (0.021)	2.063 (0.356)	10.740 (0.005)	2567
2000	0.021 (0.005)	0.151 (0.029)	0.079 (0.017)	0.298 (0.040)	-0.241* (0.086)	0.385 (0.079)	0.070 (0.021)	3.643 (0.162)	4.106 (0.128)	2430
2001	0.057 (0.012)	0.057 (0.045)	0.183 (0.038)	0.277 (0.019)	-0.102 (0.067)	0.722 (0.044)	0.205 (0.036)	8.610 (0.013)	5.586 (0.061)	2249
2002	0.010 (0.003)	0.317 (0.019)	0.212 (0.021)	0.273 (0.065)	0.062 (0.068)	0.348 (0.043)	0.036 (0.016)	3.413 (0.182)	13.291 (0.001)	2383
2003	0.033 (0.007)	0.258 (0.035)	0.171 (0.021)	0.273 (0.033)	-0.235* (0.070)	0.428 (0.072)	0.122 (0.029)	0.480 (0.787)	8.798 (0.012)	2248
2004	0.027 (0.007)	0.145 (0.036)	0.088 (0.024)	0.242 (0.031)	-0.236* (0.076)	0.353 (0.079)	0.111 (0.028)	4.720 (0.094)	9.314 (0.009)	2076
Average	0.029 (0.005)	0.168 (0.032)	0.130 (0.018)	0.262 (0.008)	-0.176 (0.042)	0.471 (0.049)	0.110 (0.019)			

Calculations are based on a sample of nonfinancial firms from Global Vantage from 1997 to 2004. GMM estimates are from the fourth-order estimator in Erickson and Whited (2000). The dependent variable is the change in the stock of cash divided by total assets.  $CF$  stands for cash flow divided by total assets;  $q$  stands for the market-to-book ratio;  $r^2$  is the measurement-error consistent estimate of the regression coefficient of determination; and  $\tau^2$  is the coefficient of determination of the measurement equation. Asymptotic standard errors are in parentheses below the yearly estimates. Fama-MacBeth standard errors are below the average estimates. An asterisk by a cash-flow coefficient indicates that its t-statistic exceeds the bootstrapped critical value.

Table 3: Cash Regressions: Canada

Year	OLS			GMM						
	$q$	$CF$	$R^2$	$q$	$CF$	$r^2$	$\tau^2$	$J$ -Test	ID Test	Obs.
1997	0.080 (0.018)	-0.098 (0.066)	0.224 (0.084)	0.249 (0.041)	-0.166* (0.088)	0.693 (0.073)	0.401 (0.063)	4.285 (0.117)	4.017 (0.134)	357
1998	0.018 (0.011)	0.211 (0.065)	0.069 (0.032)	0.331 (0.134)	0.273 (0.182)	0.294 (0.124)	0.050 (0.032)	4.701 (0.095)	7.103 (0.029)	358
1999	0.018 (0.011)	0.275 (0.070)	0.130 (0.043)	0.444 (0.237)	-0.392* (0.192)	0.418 (0.202)	0.072 (0.019)	3.161 (0.206)	3.851 (0.146)	364
2000	0.045 (0.017)	0.223 (0.072)	0.138 (0.065)	0.335 (0.100)	-0.395* (0.226)	0.532 (0.105)	0.090 (0.046)	3.189 (0.203)	7.078 (0.029)	367
2001	0.058 (0.015)	0.161 (0.047)	0.256 (0.084)	0.171 (0.015)	0.115 (0.067)	0.651 (0.109)	0.268 (0.087)	2.130 (0.345)	3.959 (0.138)	337
2002	0.023 (0.020)	0.072 (0.057)	0.045 (0.049)	0.212 (0.051)	0.072 (0.074)	0.315 (0.186)	0.242 (0.077)	3.087 (0.214)	4.892 (0.087)	328
2003	0.060 (0.022)	0.010 (0.090)	0.136 (0.070)	0.293 (0.078)	-0.164 (0.145)	0.651 (0.134)	0.219 (0.048)	1.204 (0.548)	3.685 (0.158)	319
2004	0.032 (0.010)	0.142 (0.068)	0.087 (0.032)	0.283 (0.098)	0.198 (0.122)	0.549 (0.137)	0.112 (0.043)	0.071 (0.965)	3.417 (0.181)	297
Average	0.042 (0.008)	0.125 (0.044)	0.136 (0.026)	0.290 (0.030)	-0.057 (0.092)	0.513 (0.055)	0.182 (0.043)			

Calculations are based on a sample of nonfinancial firms from Global Vantage from 1997 to 2004. GMM estimates are from the fourth-order estimator in Erickson and Whited (2000). The dependent variable is the change in the stock of cash divided by total assets.  $CF$  stands for cash flow divided by total assets;  $q$  stands for the market-to-book ratio;  $r^2$  is the measurement-error consistent estimate of the regression coefficient of determination; and  $\tau^2$  is the coefficient of determination of the measurement equation. Asymptotic standard errors are in parentheses below the yearly estimates. Fama-MacBeth standard errors are below the average estimates. An asterisk by a cash-flow coefficient indicates that its t-statistic exceeds the bootstrapped critical value.

Table 4: Cash Regressions: United Kingdom

Year	OLS			GMM						
	$q$	$CF$	$R^2$	$q$	$CF$	$r^2$	$\tau^2$	$J$ -Test	ID Test	Obs.
1997	0.013 (0.006)	0.130 (0.043)	0.065 (0.026)	0.184 (0.105)	-0.530* (0.416)	0.238 (0.078)	0.003 (0.013)	2.022 (0.364)	2.185 (0.335)	616
1998	0.004 (0.005)	0.162 (0.047)	0.038 (0.017)	0.511 (0.285)	-1.558* (1.000)	0.154 (0.163)	0.065 (0.022)	1.586 (0.452)	3.635 (0.162)	822
1999	0.016 (0.008)	0.163 (0.055)	0.069 (0.025)	0.226 (0.045)	-0.538* (0.185)	0.255 (0.089)	0.149 (0.031)	6.480 (0.039)	4.832 (0.089)	784
2000	0.034 (0.011)	-0.003 (0.064)	0.079 (0.036)	0.257 (0.036)	-0.689* (0.156)	0.540 (0.094)	0.175 (0.038)	1.653 (0.438)	5.849 (0.054)	711
2001	0.011 (0.006)	0.110 (0.044)	0.043 (0.017)	0.173 (0.034)	-0.204* (0.094)	0.250 (0.115)	0.110 (0.026)	3.780 (0.151)	5.160 (0.076)	714
2002	-0.000 (0.005)	0.195 (0.039)	0.074 (0.027)	-0.473 (0.507)	1.001 (0.870)	0.023 (0.200)	0.021 (0.026)	5.576 (0.062)	4.229 (0.121)	733
2003	0.015 (0.012)	0.223 (0.068)	0.077 (0.021)	0.230 (0.077)	-0.280* (0.198)	0.163 (0.084)	0.004 (0.009)	1.093 (0.579)	4.177 (0.124)	698
2004	-0.017 (0.007)	0.306 (0.060)	0.096 (0.035)	-0.242 (0.116)	0.842 (0.287)	0.214 (0.084)	0.105 (0.051)	2.407 (0.300)	2.185 (0.335)	613
Average	0.010 (0.005)	0.161 (0.032)	0.068 (0.007)	0.227 (0.081)	-0.363 (0.241)	0.230 (0.052)	0.079 (0.023)			

Calculations are based on a sample of nonfinancial firms from Global Vantage from 1997 to 2004. GMM estimates are from the fourth-order estimator in Erickson and Whited (2000). The dependent variable is the change in the stock of cash divided by total assets.  $CF$  stands for cash flow divided by total assets;  $q$  stands for the market-to-book ratio;  $r^2$  is the measurement-error consistent estimate of the regression coefficient of determination; and  $\tau^2$  is the coefficient of determination of the measurement equation. Asymptotic standard errors are in parentheses below the yearly estimates. Fama-MacBeth standard errors are below the average estimates. An asterisk by a cash-flow coefficient indicates that its t-statistic exceeds the bootstrapped critical value.

Table 5: Cash Regressions: Japan

Year	OLS			GMM						
	$q$	$CF$	$R^2$	$q$	$CF$	$r^2$	$\tau^2$	$J$ -Test	ID Test	Obs.
1997	0.021 (0.006)	0.283 (0.035)	0.058 (0.013)	0.229 (0.036)	-0.096 (0.091)	0.178 (0.044)	0.107 (0.029)	2.038 (0.361)	9.237 (0.010)	1984
1998	0.014 (0.004)	0.138 (0.033)	0.024 (0.008)	0.280 (0.098)	-0.446* (0.231)	0.187 (0.048)	0.054 (0.022)	2.659 (0.265)	0.546 (0.761)	2101
1999	0.009 (0.004)	0.133 (0.030)	0.018 (0.007)	0.217 (0.120)	-0.287* (0.249)	0.069 (0.032)	0.047 (0.027)	1.904 (0.386)	17.986 (0.000)	2220
2000	0.019 (0.006)	0.287 (0.033)	0.079 (0.019)	0.288 (0.092)	-0.262* (0.205)	0.214 (0.051)	0.033 (0.018)	4.080 (0.130)	7.532 (0.023)	2231
2001	0.036 (0.014)	0.227 (0.032)	0.074 (0.029)	0.238 (0.025)	-0.224* (0.078)	0.252 (0.053)	0.151 (0.037)	1.889 (0.389)	8.813 (0.012)	2345
2002	0.014 (0.004)	0.183 (0.026)	0.041 (0.010)	0.228 (0.077)	-0.187* (0.148)	0.146 (0.024)	0.127 (0.037)	11.240 (0.004)	14.975 (0.001)	2353
2003	0.017 (0.005)	0.235 (0.030)	0.059 (0.011)	0.397 (0.125)	-0.547* (0.287)	0.306 (0.051)	0.080 (0.017)	12.930 (0.002)	10.520 (0.005)	2423
2004	0.031 (0.007)	0.199 (0.034)	0.070 (0.016)	0.232 (0.054)	-0.237* (0.134)	0.278 (0.043)	0.108 (0.025)	4.532 (0.104)	4.882 (0.087)	2402
Average	0.020 (0.003)	0.211 (0.021)	0.053 (0.008)	0.264 (0.021)	-0.286 (0.051)	0.204 (0.027)	0.088 (0.015)			

Calculations are based on a sample of nonfinancial firms from Global Vantage from 1997 to 2004. GMM estimates are from the fourth-order estimator in Erickson and Whited (2000). The dependent variable is the change in the stock of cash divided by total assets.  $CF$  stands for cash flow divided by total assets;  $q$  stands for the market-to-book ratio;  $r^2$  is the measurement-error consistent estimate of the regression coefficient of determination; and  $\tau^2$  is the coefficient of determination of the measurement equation. Asymptotic standard errors are in parentheses below the yearly estimates. Fama-MacBeth standard errors are below the average estimates. An asterisk by a cash-flow coefficient indicates that its t-statistic exceeds the bootstrapped critical value.

Table 6: Cash Regressions: France

Year	OLS			GMM						
	$q$	$CF$	$R^2$	$q$	$CF$	$r^2$	$\tau^2$	$J$ -Test	ID Test	Obs.
1997	0.029 (0.010)	0.068 (0.071)	0.103 (0.065)	0.098 (0.073)	-0.390* (0.130)	0.194 (0.132)	0.292 (0.292)	0.212 (0.899)	2.286 (0.319)	233
1998	0.021 (0.009)	0.065 (0.076)	0.067 (0.038)	0.129 (0.056)	-0.430* (0.145)	0.265 (0.142)	0.165 (0.113)	0.433 (0.806)	8.097 (0.017)	318
1999	0.043 (0.027)	0.329 (0.081)	0.155 (0.061)	0.373 (0.053)	-0.611 (0.397)	0.587 (0.225)	0.116 (0.063)	1.641 (0.440)	3.952 (0.139)	312
2000	-0.019 (0.013)	0.503 (0.089)	0.159 (0.047)	-0.190 (0.068)	0.595* (0.102)	0.116 (0.076)	0.099 (0.084)	2.375 (0.305)	6.286 (0.043)	308
2001	0.038 (0.013)	0.784 (0.146)	0.127 (0.027)	-0.917 (2.092)	2.542 (0.605)	-0.433 (0.927)	-0.041 (0.101)	1.416 (0.493)	1.036 (0.596)	447
2002	0.021 (0.007)	0.154 (0.049)	0.060 (0.020)	1.246 (2.880)	-3.145 (3.281)	1.286 (3.140)	0.017 (0.037)	0.390 (0.823)	12.808 (0.002)	445
2003	0.020 (0.009)	0.161 (0.053)	0.067 (0.032)	0.137 (0.159)	-0.169 (0.159)	0.099 (0.095)	0.146 (0.219)	9.223 (0.010)	26.820 (0.000)	407
2004	0.011 (0.007)	0.173 (0.054)	0.059 (0.030)	0.661 (1.486)	-1.655 (2.423)	0.416 (0.812)	0.017 (0.046)	2.154 (0.341)	4.925 (0.085)	318
Average	0.021 (0.007)	0.280 (0.089)	0.100 (0.015)	0.192 (0.222)	-0.408 (0.580)	0.316 (0.174)	0.112 (0.033)			

Calculations are based on a sample of nonfinancial firms from Global Vantage from 1997 to 2004. GMM estimates are from the fourth-order estimator in Erickson and Whited (2000). The dependent variable is the change in the stock of cash divided by total assets.  $CF$  stands for cash flow divided by total assets;  $q$  stands for the market-to-book ratio;  $r^2$  is the measurement-error consistent estimate of the regression coefficient of determination; and  $\tau^2$  is the coefficient of determination of the measurement equation. Asymptotic standard errors are in parentheses below the yearly estimates. Fama-MacBeth standard errors are below the average estimates. An asterisk by a cash-flow coefficient indicates that its t-statistic exceeds the bootstrapped critical value.

Table 7: Cash Regressions: Germany

Year	OLS			GMM						
	$q$	$CF$	$R^2$	$q$	$CF$	$r^2$	$\tau^2$	$J$ -Test	ID Test	Obs.
1997	-0.019 (0.009)	0.051 (0.060)	0.029 (0.025)	-0.101 (0.030)	0.266 (0.144)	0.218 (0.107)	0.191 (0.102)	4.615 (0.099)	3.466 (0.177)	208
1998	0.066 (0.037)	-0.025 (0.118)	0.139 (0.115)	0.267 (0.013)	-0.418* (0.238)	0.430 (0.137)	0.248 (0.132)	2.000 (0.368)	5.818 (0.055)	270
1999	0.042 (0.022)	0.103 (0.075)	0.139 (0.117)	0.206 (0.021)	-0.182* (0.114)	0.180 (0.159)	0.206 (0.106)	0.334 (0.846)	5.636 (0.060)	279
2000	0.016 (0.008)	0.161 (0.055)	0.053 (0.021)	0.335 (0.285)	-0.721* (0.703)	0.336 (0.192)	0.048 (0.055)	1.618 (0.445)	0.822 (0.663)	301
2001	0.019 (0.012)	0.252 (0.080)	0.067 (0.029)	0.627 (0.613)	-0.244* (0.216)	0.367 (0.159)	0.030 (0.045)	2.609 (0.271)	4.565 (0.102)	330
2002	0.005 (0.010)	0.446 (0.087)	0.099 (0.034)	0.183 (0.484)	-0.157 (0.267)	0.581 (0.220)	0.026 (0.037)	5.244 (0.073)	4.177 (0.124)	384
2003	0.017 (0.009)	0.293 (0.048)	0.120 (0.037)	0.476 (1.527)	-0.237 (0.319)	0.483 (0.284)	0.035 (0.124)	2.113 (0.348)	4.690 (0.096)	367
2004	0.030 (0.018)	0.358 (0.101)	0.133 (0.051)	0.593 (0.305)	-0.084 (0.222)	0.400 (0.116)	0.051 (0.053)	0.824 (0.662)	2.427 (0.297)	214
Average	0.022 (0.009)	0.205 (0.057)	0.097 (0.015)	0.323 (0.085)	-0.222 (0.099)	0.374 (0.047)	0.104 (0.033)			

Calculations are based on a sample of nonfinancial firms from Global Vantage from 1997 to 2004. GMM estimates are from the fourth-order estimator in Erickson and Whited (2000). The dependent variable is the change in the stock of cash divided by total assets.  $CF$  stands for cash flow divided by total assets;  $q$  stands for the market-to-book ratio;  $r^2$  is the measurement-error consistent estimate of the regression coefficient of determination; and  $\tau^2$  is the coefficient of determination of the measurement equation. Asymptotic standard errors are in parentheses below the yearly estimates. Fama-MacBeth standard errors are below the average estimates. An asterisk by a cash-flow coefficient indicates that its t-statistic exceeds the bootstrapped critical value.

Table 8: Cash Regressions: USA, Small versus Large Firms

Small										
Year	OLS			GMM				$J$ -Test	ID Test	Obs.
	$q$	$CF$	$R^2$	$q$	$CF$	$\tau^2$	$\tau^2$			
1997	0.047 (0.009)	0.088 (0.040)	0.132 (0.032)	0.276 (0.020)	-0.117 (0.073)	0.610 (0.049)	0.168 (0.039)	2.839 (0.242)	7.596 (0.022)	1422
1998	0.039 (0.009)	0.185 (0.044)	0.137 (0.036)	0.236 (0.025)	-0.050 (0.087)	0.632 (0.065)	0.165 (0.046)	3.541 (0.170)	7.664 (0.022)	1397
1999	0.018 (0.005)	0.245 (0.038)	0.124 (0.027)	0.237 (0.038)	-0.117 (0.115)	0.473 (0.061)	0.077 (0.027)	2.827 (0.243)	10.364 (0.006)	1335
2000	0.027 (0.007)	0.185 (0.036)	0.101 (0.025)	0.309 (0.044)	-0.047 (0.091)	0.440 (0.077)	0.088 (0.028)	0.354 (0.838)	7.730 (0.021)	1208
2001	0.077 (0.016)	0.106 (0.050)	0.244 (0.053)	0.290 (0.018)	0.068 (0.076)	0.763 (0.042)	0.264 (0.049)	2.410 (0.300)	6.784 (0.034)	1048
2002	0.013 (0.005)	0.373 (0.024)	0.240 (0.026)	0.299 (0.087)	0.233* (0.083)	0.350 (0.051)	0.045 (0.023)	3.822 (0.148)	8.708 (0.013)	1142
2003	0.055 (0.011)	0.300 (0.042)	0.236 (0.031)	0.274 (0.035)	-0.030 (0.074)	0.439 (0.071)	0.201 (0.045)	1.045 (0.593)	14.582 (0.001)	1044
2004	0.028 (0.006)	0.176 (0.045)	0.087 (0.022)	0.210 (0.044)	-0.203* (0.153)	0.486 (0.056)	0.132 (0.045)	2.063 (0.356)	15.297 (0.000)	913
Average	0.038 (0.007)	0.207 (0.034)	0.163 (0.023)	0.266 (0.012)	-0.033 (0.047)	0.524 (0.047)	0.143 (0.025)			

Large										
Year	OLS			GMM				$J$ -Test	ID Test	Obs.
	$q$	$CF$	$R^2$	$q$	$CF$	$\tau^2$	$\tau^2$			
1997	0.002 (0.004)	0.098 (0.039)	0.023 (0.012)	0.341 (0.330)	-1.048 (1.155)	0.111 (0.109)	0.017 (0.013)	3.368 (0.186)	4.335 (0.114)	1218
1998	0.004 (0.005)	0.054 (0.028)	0.015 (0.012)	0.175 (0.069)	-0.523* (0.249)	0.129 (0.099)	0.056 (0.036)	4.386 (0.112)	4.845 (0.089)	1232
1999	0.012 (0.004)	0.044 (0.045)	0.036 (0.013)	0.136 (0.040)	-0.532* (0.201)	0.226 (0.082)	0.132 (0.057)	3.379 (0.185)	6.260 (0.044)	1222
2000	0.018 (0.005)	0.070 (0.071)	0.071 (0.024)	0.232 (0.039)	-0.762* (0.186)	0.591 (0.101)	0.075 (0.015)	7.018 (0.030)	6.034 (0.049)	1201
2001	0.008 (0.003)	0.149 (0.030)	0.088 (0.021)	0.234 (0.132)	-0.484 (0.409)	0.395 (0.130)	0.031 (0.017)	1.908 (0.385)	9.680 (0.008)	1214
2002	0.011 (0.003)	0.139 (0.026)	0.110 (0.026)	0.120 (0.031)	-0.177* (0.101)	0.307 (0.089)	0.172 (0.056)	6.255 (0.044)	0.870 (0.647)	1241
2003	0.010 (0.004)	0.135 (0.034)	0.067 (0.021)	0.244 (0.140)	-0.964* (0.680)	0.338 (0.123)	0.061 (0.019)	4.833 (0.089)	4.811 (0.090)	1204
2004	0.031 (0.017)	0.022 (0.044)	0.105 (0.060)	0.261 (0.040)	-1.043* (0.222)	0.686 (0.189)	0.157 (0.064)	3.896 (0.143)	8.579 (0.014)	1163
Average	0.012 (0.003)	0.089 (0.017)	0.064 (0.013)	0.218 (0.025)	-0.692 (0.111)	0.348 (0.073)	0.088 (0.021)			

Calculations are based on a sample of nonfinancial firms from Global Vantage from 1997 to 2004. GMM estimates are from the fourth-order estimator in Erickson and Whited (2000). The dependent variable is the change in the stock of cash divided by total assets.  $CF$  stands for cash flow divided by total assets;  $q$  stands for the market-to-book ratio;  $\tau^2$  is the measurement-error consistent estimate of the regression coefficient of determination; and  $\tau^2$  is the coefficient of determination of the measurement equation. Asymptotic standard errors are in parentheses below the yearly estimates. Fama-MacBeth standard errors are below the average estimates. An asterisk by a cash-flow coefficient indicates that its t-statistic exceeds the bootstrapped critical value.

Table 9: Cash Regressions: Japan, Small versus Large Firms

Small										
Year	OLS			GMM				$J$ -Test	ID Test	Obs.
	$q$	$CF$	$R^2$	$q$	$CF$	$\tau^2$	$\tau^2$			
1997	0.023 (0.008)	0.345 (0.060)	0.075 (0.022)	0.215 (0.020)	-0.246* (0.157)	0.231 (0.066)	0.114 (0.034)	0.509 (0.775)	8.958 (0.011)	985
1998	0.011 (0.005)	0.220 (0.052)	0.034 (0.013)	0.305 (0.128)	-0.777 (0.770)	0.209 (0.115)	0.044 (0.028)	2.979 (0.225)	4.855 (0.088)	1029
1999	0.028 (0.008)	0.290 (0.047)	0.098 (0.033)	0.115 (0.012)	-0.062 (0.155)	0.190 (0.048)	0.247 (0.069)	0.226 (0.893)	5.198 (0.074)	1119
2000	0.049 (0.024)	0.204 (0.045)	0.096 (0.058)	0.261 (0.003)	-0.262* (0.075)	0.295 (0.130)	0.212 (0.044)	5.932 (0.052)	5.371 (0.068)	1174
2001	0.016 (0.006)	0.187 (0.040)	0.044 (0.015)	0.147 (0.020)	-0.299 (0.237)	0.184 (0.047)	0.159 (0.044)	0.500 (0.779)	5.961 (0.051)	1177
2002	0.021 (0.008)	0.313 (0.047)	0.074 (0.018)	0.307 (0.029)	-0.345 (0.252)	0.243 (0.070)	0.093 (0.026)	7.295 (0.026)	8.190 (0.017)	1245
2003	0.025 (0.006)	0.262 (0.050)	0.085 (0.021)	0.097 (0.014)	-0.215 (0.227)	0.220 (0.045)	0.299 (0.067)	8.750 (0.013)	7.296 (0.026)	1295
2004	0.044 (0.009)	0.206 (0.051)	0.113 (0.028)	0.264 (0.037)	-0.218 (0.141)	0.286 (0.060)	0.155 (0.035)	2.202 (0.332)	5.525 (0.063)	1289
Average	0.027 (0.005)	0.253 (0.021)	0.077 (0.010)	0.214 (0.030)	-0.303 (0.074)	0.232 (0.014)	0.165 (0.030)			

Large										
Year	OLS			GMM				$J$ -Test	ID Test	Obs.
	$q$	$CF$	$R^2$	$q$	$CF$	$\tau^2$	$\tau^2$			
1997	0.014 (0.008)	0.257 (0.046)	0.041 (0.012)	0.262 (0.117)	-0.143 (0.211)	0.093 (0.056)	0.012 (0.013)	2.176 (0.337)	8.300 (0.016)	999
1998	0.017 (0.006)	0.104 (0.043)	0.022 (0.010)	0.246 (0.090)	-0.386* (0.212)	0.160 (0.054)	0.042 (0.027)	1.095 (0.578)	7.790 (0.020)	1072
1999	0.009 (0.006)	0.102 (0.042)	0.013 (0.009)	0.171 (0.082)	-0.215* (0.176)	0.050 (0.034)	0.180 (0.157)	3.804 (0.149)	10.292 (0.006)	1113
2000	0.011 (0.009)	0.300 (0.047)	0.074 (0.022)	0.529 (0.334)	-0.731* (0.670)	0.192 (0.092)	0.063 (0.025)	2.840 (0.242)	9.568 (0.008)	1112
2001	0.023 (0.010)	0.275 (0.046)	0.066 (0.021)	0.157 (0.044)	-0.015 (0.112)	0.098 (0.048)	0.145 (0.052)	0.761 (0.684)	9.881 (0.007)	1171
2002	0.010 (0.005)	0.192 (0.032)	0.043 (0.014)	0.195 (0.117)	-0.069 (0.186)	0.118 (0.036)	0.153 (0.081)	3.825 (0.148)	9.947 (0.007)	1176
2003	0.011 (0.006)	0.155 (0.037)	0.041 (0.013)	0.766 (0.988)	-1.403 (2.085)	0.510 (0.425)	0.066 (0.035)	4.927 (0.085)	9.369 (0.009)	1178
2004	0.012 (0.006)	0.161 (0.039)	0.038 (0.015)	0.364 (0.283)	-0.655 (0.670)	0.240 (0.117)	0.024 (0.023)	2.625 (0.269)	9.765 (0.008)	1164
Average	0.013 (0.002)	0.193 (0.027)	0.042 (0.007)	0.336 (0.075)	-0.452 (0.165)	0.183 (0.051)	0.086 (0.023)			

Calculations are based on a sample of nonfinancial firms from Global Vantage from 1997 to 2004. GMM estimates are from the fourth-order estimator in Erickson and Whited (2000). The dependent variable is the change in the stock of cash divided by total assets.  $CF$  stands for cash flow divided by total assets;  $q$  stands for the market-to-book ratio;  $\tau^2$  is the measurement-error consistent estimate of the regression coefficient of determination; and  $\tau^2$  is the coefficient of determination of the measurement equation. Asymptotic standard errors are in parentheses below the yearly estimates. Fama-MacBeth standard errors are below the average estimates. An asterisk by a cash-flow coefficient indicates that its t-statistic exceeds the bootstrapped critical value.

Table 10: Cash Regressions: USA, High versus Low  $\rho$  Firms

Low $\rho$										
Year	OLS			GMM						
	$q$	$CF$	$R^2$	$q$	$CF$	$r^2$	$\tau^2$	$J$ -Test	ID Test	Obs.
1997	0.019 (0.009)	0.126 (0.033)	0.056 (0.024)	0.255 (0.008)	-0.020 (0.077)	0.618 (0.049)	0.131 (0.027)	9.925 (0.007)	4.909 (0.086)	1059
1998	0.046 (0.016)	0.152 (0.060)	0.135 (0.056)	0.224 (0.014)	0.087 (0.084)	0.657 (0.088)	0.225 (0.049)	3.549 (0.170)	5.727 (0.057)	994
1999	0.019 (0.009)	0.178 (0.046)	0.082 (0.028)	0.286 (0.018)	0.060 (0.092)	0.583 (0.060)	0.098 (0.021)	2.407 (0.300)	9.495 (0.009)	888
2000	0.027 (0.014)	0.226 (0.062)	0.090 (0.036)	0.330 (0.010)	0.219 (0.133)	0.674 (0.054)	0.123 (0.027)	2.802 (0.246)	10.992 (0.004)	728
2001	0.080 (0.031)	0.151 (0.067)	0.240 (0.100)	0.307 (0.005)	0.241 (0.120)	0.780 (0.047)	0.204 (0.056)	1.182 (0.554)	4.311 (0.116)	634
2002	0.026 (0.008)	0.313 (0.031)	0.233 (0.034)	0.207 (0.021)	0.301* (0.053)	0.467 (0.066)	0.146 (0.045)	8.069 (0.018)	6.659 (0.036)	736
2003	0.061 (0.015)	0.290 (0.058)	0.228 (0.044)	0.082 (0.043)	0.272* (0.070)	0.141 (0.079)	0.380 (0.214)	5.555 (0.062)	8.763 (0.013)	666
2004	0.029 (0.007)	0.155 (0.066)	0.079 (0.027)	0.405 (0.094)	0.085 (0.150)	0.471 (0.080)	0.055 (0.027)	3.866 (0.145)	4.005 (0.135)	585
Average	0.038 (0.008)	0.199 (0.025)	0.143 (0.028)	0.262 (0.034)	0.156 (0.041)	0.549 (0.069)	0.170 (0.036)			
High $\rho$										
Year	OLS			GMM						
	$q$	$CF$	$R^2$	$q$	$CF$	$r^2$	$\tau^2$	$J$ -Test	ID Test	Obs.
1997	0.063 (0.015)	-0.144 (0.121)	0.212 (0.066)	0.234 (0.013)	-0.602* (0.176)	0.761 (0.059)	0.262 (0.067)	0.337 (0.845)	7.057 (0.029)	766
1998	0.022 (0.012)	0.208 (0.074)	0.106 (0.053)	0.170 (0.006)	-0.296* (0.126)	0.535 (0.080)	0.184 (0.053)	7.243 (0.027)	4.083 (0.130)	824
1999	0.015 (0.006)	0.198 (0.055)	0.105 (0.041)	0.162 (0.012)	-0.270* (0.088)	0.445 (0.065)	0.122 (0.024)	4.502 (0.105)	6.690 (0.035)	852
2000	0.013 (0.005)	0.131 (0.046)	0.064 (0.021)	0.238 (0.045)	-0.446* (0.193)	0.495 (0.062)	0.070 (0.023)	1.538 (0.464)	4.286 (0.117)	863
2001	0.019 (0.005)	0.074 (0.045)	0.078 (0.024)	0.184 (0.020)	-0.260 (0.094)	0.526 (0.066)	0.107 (0.023)	1.640 (0.440)	8.696 (0.013)	820
2002	0.010 (0.004)	0.252 (0.039)	0.157 (0.031)	0.243 (0.054)	-0.330 (0.190)	0.454 (0.087)	0.047 (0.017)	2.503 (0.286)	10.916 (0.004)	803
2003	0.022 (0.012)	0.191 (0.073)	0.119 (0.028)	0.260 (0.017)	-0.584* (0.131)	0.562 (0.101)	0.107 (0.035)	2.095 (0.351)	7.229 (0.027)	765
2004	0.028 (0.017)	0.155 (0.056)	0.102 (0.051)	0.303 (0.013)	-0.930* (0.165)	0.522 (0.116)	0.083 (0.036)	0.050 (0.975)	5.006 (0.082)	716
Average	0.024 (0.006)	0.133 (0.044)	0.118 (0.017)	0.224 (0.017)	-0.465 (0.082)	0.538 (0.035)	0.123 (0.025)			

Calculations are based on a sample of nonfinancial firms from Global Vantage from 1997 to 2004. GMM estimates are from the fourth-order estimator in Erickson and Whited (2000). The sample is divided by  $\rho$ , an estimate of the serial correlation of the shock to profits. The dependent variable is the change in the stock of cash divided by total assets.  $CF$  stands for cash flow divided by total assets;  $q$  stands for the market-to-book ratio;  $r^2$  is the measurement-error consistent estimate of the regression coefficient of determination; and  $\tau^2$  is the coefficient of determination of the measurement equation. Asymptotic standard errors are in parentheses below the yearly estimates. Fama-MacBeth standard errors are below the average estimates. An asterisk by a cash-flow coefficient indicates that its t-statistic exceeds the bootstrapped critical value.

Table 11: Cash Regressions: Japan, High versus Low  $\rho$  Firms

Low $\rho$										
Year	OLS			GMM						
	$q$	$CF$	$R^2$	$q$	$CF$	$r^2$	$\tau^2$	$J$ -Test	ID Test	Obs.
1997	0.022 (0.008)	0.190 (0.057)	0.044 (0.017)	0.136 (0.047)	-0.002 (0.112)	0.126 (0.056)	0.130 (0.067)	0.944 (0.624)	6.517 (0.038)	598
1998	0.015 (0.007)	0.178 (0.054)	0.032 (0.015)	0.210 (0.061)	-0.177 (0.145)	0.207 (0.070)	0.097 (0.041)	1.342 (0.511)	6.476 (0.039)	643
1999	0.007 (0.007)	0.249 (0.052)	0.043 (0.016)	0.399 (0.238)	-0.296 (0.387)	0.112 (0.069)	0.023 (0.024)	0.805 (0.669)	0.981 (0.612)	691
2000	0.056 (0.033)	0.279 (0.061)	0.127 (0.089)	0.254 (0.023)	-0.159* (0.054)	0.155 (0.056)	0.094 (0.029)	2.603 (0.272)	7.332 (0.026)	758
2001	0.012 (0.007)	0.242 (0.047)	0.057 (0.020)	0.123 (0.022)	0.001 (0.076)	0.143 (0.037)	0.162 (0.047)	7.228 (0.027)	8.510 (0.014)	794
2002	0.019 (0.009)	0.316 (0.061)	0.079 (0.022)	0.122 (0.052)	0.072 (0.159)	0.134 (0.058)	0.126 (0.075)	9.212 (0.010)	4.155 (0.125)	889
2003	0.025 (0.007)	0.276 (0.060)	0.088 (0.023)	0.090 (0.012)	0.112 (0.068)	0.142 (0.031)	0.366 (0.073)	7.898 (0.019)	7.210 (0.027)	924
2004	0.039 (0.011)	0.296 (0.062)	0.117 (0.032)	0.149 (0.019)	0.020 (0.085)	0.187 (0.041)	0.243 (0.068)	3.292 (0.193)	7.276 (0.026)	905
Average	0.024 (0.006)	0.253 (0.017)	0.073 (0.013)	0.185 (0.036)	-0.054 (0.050)	0.151 (0.011)	0.155 (0.037)			
High $\rho$										
Year	OLS			GMM						
	$q$	$CF$	$R^2$	$q$	$CF$	$r^2$	$\tau^2$	$J$ -Test	ID Test	Obs.
1997	0.030 (0.012)	0.317 (0.067)	0.086 (0.028)	0.227 (0.019)	-0.133 (0.113)	0.241 (0.076)	0.126 (0.052)	0.234 (0.890)	4.412 (0.110)	703
1998	0.005 (0.006)	0.089 (0.061)	0.007 (0.007)	0.166 (0.088)	-0.283 (0.218)	0.086 (0.060)	0.056 (0.043)	2.238 (0.327)	7.805 (0.020)	736
1999	0.019 (0.007)	0.156 (0.047)	0.038 (0.013)	0.148 (0.079)	-0.105 (0.176)	0.079 (0.053)	0.082 (0.062)	3.789 (0.150)	11.866 (0.003)	765
2000	0.036 (0.012)	0.386 (0.063)	0.170 (0.053)	0.120 (0.029)	0.140 (0.100)	0.134 (0.063)	0.239 (0.104)	6.980 (0.031)	10.167 (0.006)	763
2001	0.030 (0.008)	0.201 (0.054)	0.056 (0.017)	0.184 (0.070)	-0.177 (0.195)	0.117 (0.046)	0.114 (0.068)	5.395 (0.067)	10.160 (0.006)	778
2002	0.017 (0.006)	0.143 (0.045)	0.042 (0.017)	0.579 (0.968)	-0.856 (1.706)	0.213 (0.296)	0.014 (0.031)	6.406 (0.041)	5.042 (0.080)	766
2003	0.015 (0.006)	0.149 (0.045)	0.045 (0.020)	0.565 (0.587)	-1.205 (1.503)	0.405 (0.325)	0.021 (0.029)	2.688 (0.261)	10.720 (0.005)	755
2004	0.022 (0.008)	0.118 (0.048)	0.041 (0.020)	0.310 (0.149)	-0.468 (0.342)	0.272 (0.106)	0.080 (0.050)	1.695 (0.428)	5.319 (0.070)	727
Average	0.022 (0.004)	0.195 (0.037)	0.061 (0.017)	0.287 (0.065)	-0.386 (0.157)	0.193 (0.040)	0.092 (0.025)			

Calculations are based on a sample of nonfinancial firms from Global Vantage from 1997 to 2004. GMM estimates are from the fourth-order estimator in Erickson and Whited (2000). The sample is divided by  $\rho$ , an estimate of the serial correlation of the shock to profits. The dependent variable is the change in the stock of cash divided by total assets.  $CF$  stands for cash flow divided by total assets;  $q$  stands for the market-to-book ratio;  $r^2$  is the measurement-error consistent estimate of the regression coefficient of determination; and  $\tau^2$  is the coefficient of determination of the measurement equation. Asymptotic standard errors are in parentheses below the yearly estimates. Fama-MacBeth standard errors are below the average estimates. An asterisk by a cash-flow coefficient indicates that its t-statistic exceeds the bootstrapped critical value.

Table 12: Cash Regressions: USA, High versus Low  $\sigma_v$  Firms

Low $\sigma_v$										
Year	OLS			GMM						
	$q$	$CF$	$R^2$	$q$	$CF$	$r^2$	$\tau^2$	$J$ -Test	ID Test	Obs.
1997	0.018 (0.007)	0.094 (0.036)	0.053 (0.020)	0.253 (0.029)	-0.399* (0.108)	0.552 (0.065)	0.100 (0.029)	7.229 (0.027)	11.377 (0.003)	1346
1998	0.026 (0.016)	0.143 (0.041)	0.095 (0.054)	0.137 (0.042)	-0.095 (0.132)	0.109 (0.070)	0.040 (0.042)	6.451 (0.040)	6.053 (0.048)	1282
1999	0.008 (0.006)	0.136 (0.043)	0.050 (0.023)	0.295 (0.054)	-0.570* (0.211)	0.459 (0.096)	0.044 (0.017)	5.597 (0.061)	7.344 (0.025)	1160
2000	0.009 (0.006)	0.056 (0.069)	0.018 (0.013)	0.208 (0.048)	-0.447* (0.201)	0.371 (0.089)	0.074 (0.035)	6.660 (0.036)	6.759 (0.034)	1013
2001	0.012 (0.009)	0.169 (0.074)	0.065 (0.036)	0.161 (0.019)	-0.022 (0.105)	0.405 (0.162)	0.093 (0.039)	6.796 (0.033)	11.431 (0.003)	945
2002	0.003 (0.004)	0.321 (0.033)	0.241 (0.039)	0.099 (0.013)	0.217* (0.048)	0.155 (0.044)	0.097 (0.046)	7.178 (0.028)	10.786 (0.005)	990
2003	0.024 (0.011)	0.177 (0.089)	0.112 (0.030)	0.283 (0.014)	-0.590* (0.146)	0.477 (0.093)	0.087 (0.027)	1.095 (0.578)	5.214 (0.074)	929
2004	0.018 (0.007)	0.133 (0.086)	0.068 (0.024)	0.146 (0.025)	-0.185 (0.120)	0.277 (0.093)	0.117 (0.041)	6.613 (0.037)	6.205 (0.045)	843
Average	0.015 (0.003)	0.154 (0.028)	0.088 (0.024)	0.198 (0.026)	-0.261 (0.101)	0.351 (0.056)	0.082 (0.010)			

High $\sigma_v$										
Year	OLS			GMM						
	$q$	$CF$	$R^2$	$q$	$CF$	$r^2$	$\tau^2$	$J$ -Test	ID Test	Obs.
1997	0.049 (0.011)	0.064 (0.051)	0.130 (0.039)	0.276 (0.025)	-0.077 (0.088)	0.686 (0.045)	0.159 (0.048)	1.659 (0.436)	4.884 (0.087)	997
1998	0.036 (0.009)	0.159 (0.049)	0.114 (0.036)	0.307 (0.035)	-0.036 (0.101)	0.580 (0.071)	0.098 (0.029)	2.666 (0.264)	4.138 (0.126)	1096
1999	0.016 (0.005)	0.213 (0.038)	0.111 (0.028)	0.165 (0.017)	0.020 (0.051)	0.403 (0.065)	0.132 (0.023)	8.717 (0.013)	7.557 (0.023)	1152
2000	0.027 (0.007)	0.189 (0.034)	0.113 (0.024)	0.308 (0.036)	-0.145 (0.099)	0.601 (0.044)	0.088 (0.022)	0.059 (0.971)	2.465 (0.292)	1159
2001	0.076 (0.016)	0.036 (0.054)	0.248 (0.052)	0.246 (0.027)	-0.092 (0.080)	0.803 (0.055)	0.226 (0.039)	2.226 (0.329)	10.149 (0.006)	1051
2002	0.020 (0.005)	0.260 (0.030)	0.152 (0.025)	0.287 (0.025)	0.084 (0.068)	0.510 (0.051)	0.074 (0.018)	1.013 (0.602)	2.194 (0.334)	1123
2003	0.042 (0.010)	0.294 (0.039)	0.215 (0.029)	0.289 (0.019)	-0.077 (0.074)	0.567 (0.071)	0.140 (0.030)	0.091 (0.956)	9.164 (0.010)	1089
2004	0.034 (0.011)	0.163 (0.042)	0.107 (0.038)	0.285 (0.005)	-0.122 (0.090)	0.614 (0.054)	0.148 (0.028)	2.167 (0.338)	4.438 (0.109)	1011
Average	0.038 (0.007)	0.172 (0.031)	0.149 (0.019)	0.270 (0.017)	-0.056 (0.027)	0.596 (0.042)	0.133 (0.017)			

Calculations are based on a sample of nonfinancial firms from Global Vantage from 1997 to 2004. GMM estimates are from the fourth-order estimator in Erickson and Whited (2000). The sample is divided by  $\sigma_v$ , an estimate of the serial correlation of the shock to profits. The dependent variable is the change in the stock of cash divided by total assets.  $CF$  stands for cash flow divided by total assets;  $q$  stands for the market-to-book ratio;  $r^2$  is the measurement-error consistent estimate of the regression coefficient of determination; and  $\tau^2$  is the coefficient of determination of the measurement equation. Asymptotic standard errors are in parentheses below the yearly estimates. Fama-MacBeth standard errors are below the average estimates. An asterisk by a cash-flow coefficient indicates that its t-statistic exceeds the bootstrapped critical value.

Table 13: Cash Regressions: Japan, High versus Low  $\sigma_v$  Firms

Low $\sigma_v$										
Year	OLS			GMM						
	$q$	$CF$	$R^2$	$q$	$CF$	$r^2$	$\tau^2$	$J$ -Test	ID Test	Obs.
1997	0.024 (0.010)	0.295 (0.053)	0.067 (0.021)	0.231 (0.041)	-0.069 (0.114)	0.195 (0.067)	0.104 (0.053)	1.533 (0.465)	5.562 (0.062)	920
1998	0.019 (0.006)	0.142 (0.046)	0.035 (0.013)	0.227 (0.089)	-0.323 (0.218)	0.176 (0.061)	0.083 (0.050)	1.030 (0.597)	6.344 (0.042)	974
1999	0.009 (0.005)	0.175 (0.041)	0.030 (0.011)	0.225 (0.178)	-0.317 (0.417)	0.078 (0.050)	0.041 (0.050)	1.130 (0.568)	10.244 (0.006)	1047
2000	0.020 (0.009)	0.354 (0.044)	0.117 (0.031)	0.319 (0.121)	-0.352 (0.314)	0.215 (0.068)	0.063 (0.046)	5.222 (0.073)	9.081 (0.011)	1061
2001	0.021 (0.007)	0.276 (0.048)	0.065 (0.015)	0.184 (0.085)	-0.107 (0.219)	0.104 (0.040)	0.114 (0.077)	4.174 (0.124)	8.882 (0.012)	1086
2002	0.011 (0.005)	0.184 (0.035)	0.041 (0.013)	0.225 (0.124)	-0.214 (0.246)	0.127 (0.046)	0.049 (0.046)	4.533 (0.104)	9.349 (0.009)	1054
2003	0.015 (0.006)	0.205 (0.042)	0.058 (0.019)	0.194 (0.078)	-0.153 (0.181)	0.143 (0.042)	0.080 (0.058)	5.238 (0.073)	8.643 (0.013)	1055
2004	0.022 (0.007)	0.207 (0.047)	0.054 (0.016)	0.466 (0.426)	-0.601 (0.799)	0.337 (0.231)	0.047 (0.055)	2.861 (0.239)	10.177 (0.006)	984
Average	0.018 (0.002)	0.230 (0.025)	0.058 (0.010)	0.259 (0.033)	-0.267 (0.061)	0.172 (0.029)	0.073 (0.010)			

High $\sigma_v$										
Year	OLS			GMM						
	$q$	$CF$	$R^2$	$q$	$CF$	$r^2$	$\tau^2$	$J$ -Test	ID Test	Obs.
1997	0.017 (0.008)	0.234 (0.048)	0.038 (0.015)	0.262 (0.095)	-0.182 (0.197)	0.144 (0.071)	0.057 (0.038)	0.770 (0.681)	7.466 (0.024)	862
1998	0.008 (0.005)	0.116 (0.051)	0.011 (0.007)	0.325 (0.177)	-0.447 (0.341)	0.161 (0.090)	0.032 (0.026)	0.944 (0.624)	11.559 (0.003)	909
1999	0.003 (0.007)	0.068 (0.045)	0.004 (0.005)	0.137 (0.095)	-0.113 (0.149)	0.029 (0.028)	0.041 (0.053)	1.093 (0.579)	4.251 (0.119)	944
2000	0.066 (0.032)	0.205 (0.051)	0.127 (0.094)	0.248 (0.011)	-0.123 (0.078)	0.227 (0.123)	0.158 (0.094)	2.042 (0.360)	9.374 (0.009)	1013
2001	0.017 (0.007)	0.165 (0.045)	0.037 (0.018)	0.137 (0.022)	-0.039 (0.066)	0.152 (0.040)	0.170 (0.062)	2.760 (0.252)	5.497 (0.064)	1057
2002	0.025 (0.010)	0.233 (0.050)	0.068 (0.023)	0.276 (0.018)	-0.354* (0.112)	0.416 (0.084)	0.118 (0.028)	6.243 (0.044)	8.743 (0.013)	1136
2003	0.024 (0.007)	0.233 (0.060)	0.070 (0.021)	0.107 (0.019)	-0.004 (0.087)	0.157 (0.037)	0.286 (0.075)	7.986 (0.018)	11.970 (0.003)	1174
2004	0.040 (0.011)	0.171 (0.056)	0.086 (0.032)	0.260 (0.036)	-0.389* (0.148)	0.314 (0.067)	0.150 (0.037)	3.256 (0.196)	10.626 (0.005)	1195
Average	0.025 (0.007)	0.178 (0.022)	0.055 (0.014)	0.219 (0.028)	-0.206 (0.060)	0.200 (0.042)	0.127 (0.030)			

Calculations are based on a sample of nonfinancial firms from Global Vantage from 1997 to 2004. GMM estimates are from the fourth-order estimator in Erickson and Whited (2000). The sample is divided by  $\sigma_v$ , an estimate of the serial correlation of the shock to profits. The dependent variable is the change in the stock of cash divided by total assets.  $CF$  stands for cash flow divided by total assets;  $q$  stands for the market-to-book ratio;  $r^2$  is the measurement-error consistent estimate of the regression coefficient of determination; and  $\tau^2$  is the coefficient of determination of the measurement equation. Asymptotic standard errors are in parentheses below the yearly estimates. Fama-MacBeth standard errors are below the average estimates. An asterisk by a cash-flow coefficient indicates that its t-statistic exceeds the bootstrapped critical value.

Table 14: Monte Carlo Performance of GMM and OLS Estimators

	OLS	GMM3	GMM4	GMM5	GMM6
Sample Size = 202					
$E(\hat{\alpha}_1)$	0.180	-0.234	-0.223	-0.168	-0.201
$MAD(\hat{\alpha}_1)$	0.481	0.409	0.299	0.326	0.342
$P( \hat{\alpha}_1 - \alpha_1  \leq 0.2   \alpha_1  )$	0.001	0.085	0.130	0.150	0.130
T-test Size		0.032	0.062	0.072	0.090
J-test Size		0.032	0.062	0.072	0.090
Sample Size = 1200					
$E(\hat{\alpha}_1)$	0.183	-0.349	-0.288	-0.275	-0.329
$MAD(\hat{\alpha}_1)$	0.483	0.277	0.087	0.094	0.134
$P( \hat{\alpha}_1 - \alpha_1  \leq 0.2   \alpha_1  )$	0.000	0.151	0.565	0.486	0.357
T-Test Size		0.001	0.008	0.007	0.011
J-test Size		0.032	0.062	0.072	0.090
Sample Size = 2576					
$E(\hat{\alpha}_1)$	0.183	-0.345	-0.311	-0.290	-0.300
$MAD(\hat{\alpha}_1)$	0.483	0.213	0.066	0.055	0.049
$P( \hat{\alpha}_1 - \alpha_1  \leq 0.2   \alpha_1  )$	0.000	0.202	0.668	0.675	0.761
T-Test Size		0.001	0.002	0.001	0.003
J-test Size		0.032	0.062	0.072	0.090

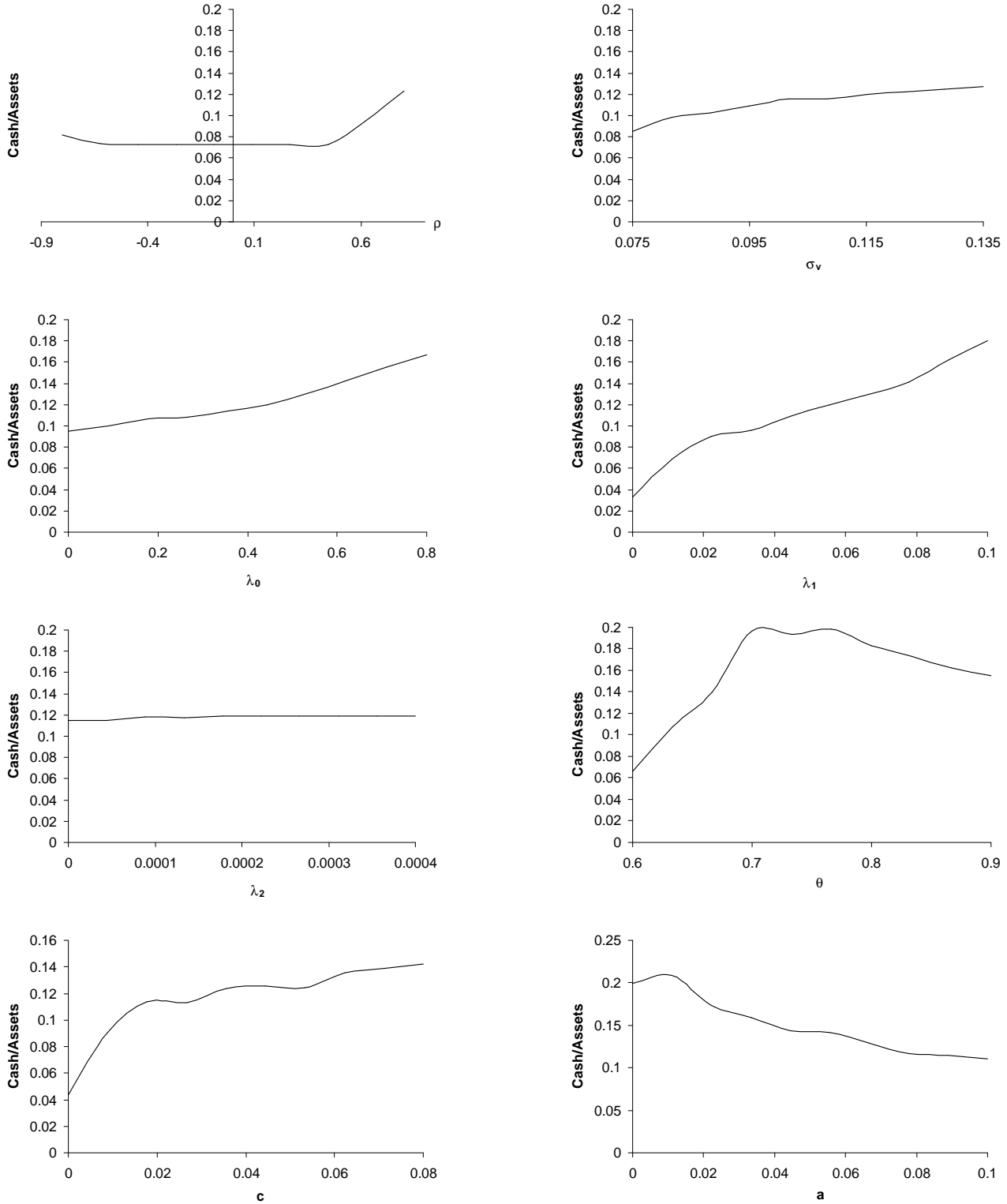
Indicated expectations and probabilities are estimates based on 10,000 Monte Carlo samples of size 336. The samples are generated by

$$\begin{aligned} y_i &= q_i\beta + w_i\alpha + u_i \\ x_i &= \gamma + \chi_i + \varepsilon_i, \end{aligned}$$

in which  $\chi_i$  and  $\varepsilon_i$  are distributed as a chi-squared variables.  $u_i$  is distributed as a negative chi-squared variable.  $GMMn$  denotes the GMM estimator based on moments up to order  $M = n$ . OLS denotes estimates obtained by regressing  $y_i$  on  $x_i$  and  $w_i$ . MAD denotes mean absolute deviation. “T-Test Size” refers to the actual size of a nominal 5% test of the null hypothesis that  $\alpha_1$  equals its true value. “J-Test Size” refers to the actual size of a nominal 10% test of the overidentifying restrictions.

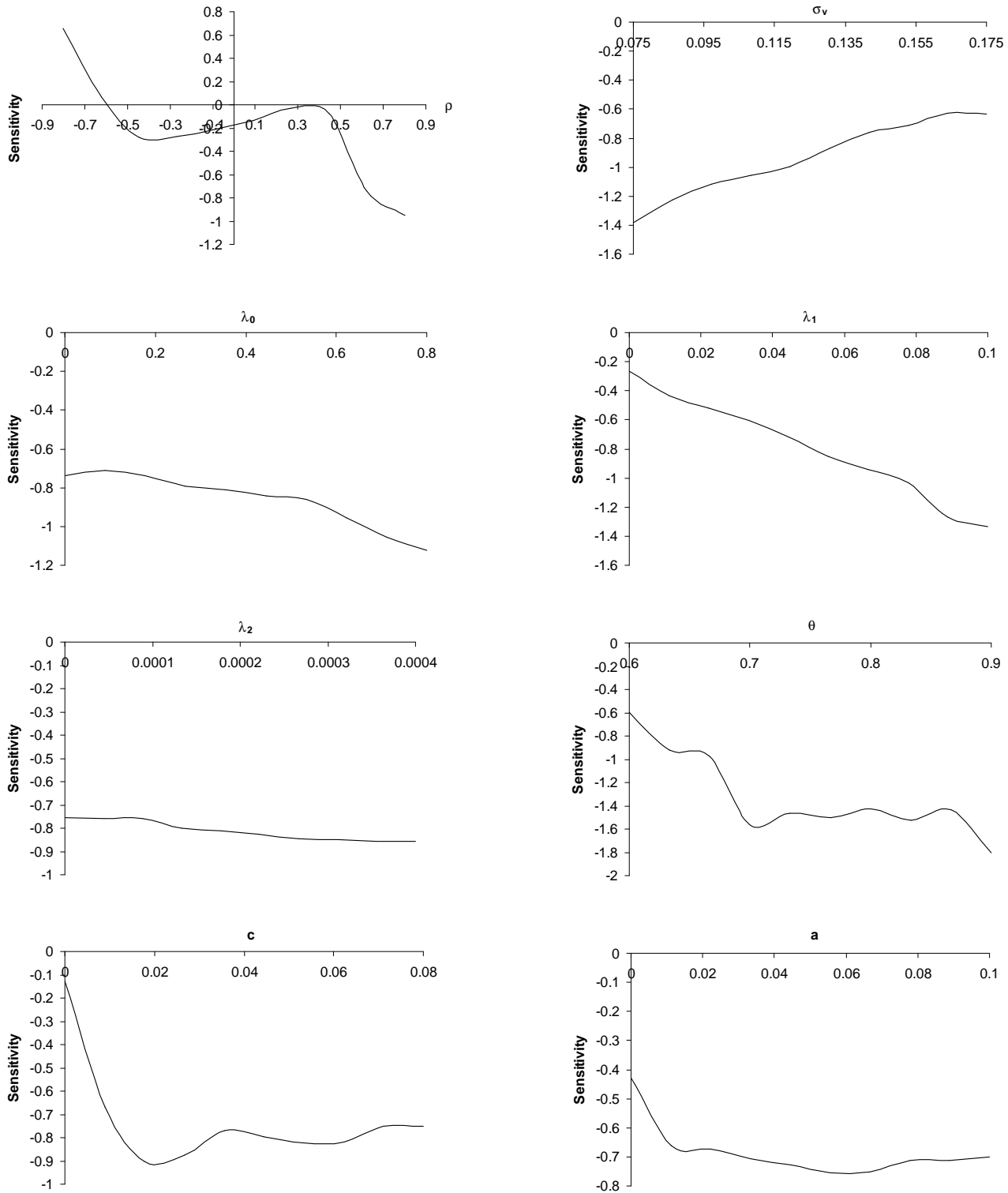
**True Value:**  $\alpha_1 = -0.3$ .

Figure 1: The Ratio of Cash to Assets



This figure depicts the relation between various model parameters and the ratio of the stock of cash to assets. The serial correlation of income shocks is  $\rho$ ; the standard deviation of the innovations to these shocks is  $\sigma_v$ ;  $\lambda_0$ ,  $\lambda_1$ , and  $\lambda_2$  are fixed, linear, and quadratic costs of external finance;  $\theta$  is the curvature of the production function; and  $c$  and  $a$  are fixed and quadratic costs of adjusting the capital stock.

Figure 2: The Cash-Flow Sensitivity of Cash



This figure depicts the relation between various model parameters and the sensitivity of the change in the cash stock to cash flow, holding constant Tobin's  $q$ . The serial correlation of income shocks is  $\rho$ ; the standard deviation of the innovations to these shocks is  $\sigma_v$ ;  $\lambda_0$ ,  $\lambda_1$ , and  $\lambda_2$  are fixed, linear, and quadratic costs of external finance;  $\theta$  is the curvature of the production function; and  $c$  and  $a$  are fixed and quadratic costs of adjusting the capital stock.