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Altunok, Fatih and Mitchell, Karlyn and Pearce, Douglas

Central Bank of Turkey, North Carolina State University

10 August 2015

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Fatih Altunok
Research Department
Central Bank of Turkey
fatih.altunok@tcmb.gov.tr
phone: +905380360444

Karlyn Mitchell
Department of Business Management
Poole College of Management
North Carolina State University
karlyn_mitchell@ncsu.edu
phone: 919 515 5593

Douglas K. Pearce
Department of Economics
Poole College of Management
North Carolina State University
doug_pearce@ncsu.edu
phone: 919 513 2880

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Abstract

We investigate whether a trade credit channel mitigates monetary policy tightenings intended to slow economic activity. Unlike prior research, we study this issue using quarterly firm-level data for nearly the universe of non-financial public corporations and using more precise measures of their credit market access. We estimate firm-level models of the supply and demand for trade credit from 1988 to 2008. Our evidence suggests that policy tightenings evoke a flow of trade credit from public firms commensurate with their credit market access which goes primarily to private firms, a previously undocumented finding.

I. Introduction

This paper reexamines the role of trade credit in the transmission mechanism of US monetary policy. In the traditional Keynesian view monetary policy impulses affect aggregate demand by influencing interest rates and investment (the interest rate channel). Empirical evidence that business investment is fairly insensitive to interest rates has led researchers to propose alternative transmission mechanisms, among them the credit channel based on frictions in credit markets (e.g., Bernanke and Gertler, 1995). Along one avenue in this channel, the bank lending channel, monetary policy is thought to affect aggregate demand by shifting the supply of bank loans, influencing the ease with which firms finance normal operations. Meltzer (1960) notes that a trade credit channel may work against this bank lending channel. He finds evidence that cash-rich firms increase trade credit to cash-poor firms in tight-money periods, filling the void left by reduced bank lending. Since Meltzer (1960), the trade credit channel has received comparatively little attention (Mateut, 2005) and its existence in the US remains controversial.[1] Neither Gertler and Gilchrist (1993) nor Oliner and Rudebusch (1996) find evidence that trade credit expands during tight-money periods while Nilsen (2002) and Choi and Kim (2005) do.[2] Garcia-Appendini and Montoriol-Garriga (2013) and Yang (2011) examine the supply and demand for trade credit around the 2008 financial crisis but their studies are focused too narrowly to provide general conclusions.

We revisit the question of a trade credit channel in the US using an innovative empirical approach. Specifically, we estimate models of the supply and demand for trade credit on quarterly firm-level data for essentially the universe of non-financial Compustat firms from 1988 through 2008.[3] In addition to including controls for the stance of monetary policy and firm-specific factors, our models include controls which better distinguish firms with good credit

[1] Brechling and Lipsey (1963) provide a theoretical basis for Meltzer's empirical results and note the potential for trade credit to contravene monetary policy. Other theoretical arguments consistent with a trade credit channel appear in Schwartz (1974), Biais and Gollier (1997), Wilner (2000), Burkart and Ellingsen (2004) and Mateut et al. (2006). Empirical evidence from the UK consistently supports the existence of a trade credit channel, including Atanasova and Wilson (2003), Mateut et al. (2006) and Guariglia and Mateut (2006). Elsewhere the empirical evidence is mixed: whereas Fishman and Love (2003) and Love et al. (2007) find evidence of a trade credit channel in multi-country studies, Marotta (1997) and De Blasio (2005) find little evidence of a trade credit channel in Italy and Cook (1999) finds no evidence in Russia.

[2] Gertler and Gilchrist (1993) and Oliner and Rudebusch (1996) study the time series behavior of aggregate accounts receivable and payable for manufacturing firms in the Quarterly Financial Reports from the mid 1970s through 1991. We describe the work of Nilsen (2002) and Choi and Kim (2005) more fully below.

[3] We end our sample period after 2008 to prevent distortions attributable to the financial crisis. The empirical results we report are qualitatively unchanged if we drop the last two quarters of 2008.

market access able to supply trade credit from firms with poor credit market access having need of trade credit.

To preview our results, we find evidence consistent with a trade credit channel involving both public and private firms. Specifically, we find that in tight-money periods public firms with good credit market access expand their accounts receivable (i.e., supply trade credit) commensurate with their credit market access as envisioned by Meltzer (1960). We also find that tight money leads public firms to expand their accounts payable (i.e., demand for trade credit) negligibly irrespective of their credit market access. These findings are both novel to the literature and surprising because they imply that some demand for trade credit is missing because one firm's receivables are another firm's payables. Since our sample covers nearly the universe of public corporations, we conclude that the missing trade credit demand comes from private firms not in our sample but which have poor credit market access, a conclusion for which we find support in the literature.

Our results are significant for at least two reasons. First, they suggest that a greater tightening of monetary policy is needed to achieve a desired degree of restraint than in the absence of the trade credit channel. Second, they suggest that the effects of tighter monetary policy fall unevenly on public and private firms.

Our paper relates directly to work by Nilsen (2002) and Choi and Kim (2005), whose evidence on the trade credit channel complements and contrasts with ours. Nilsen uses VAR models to study how a reduction in bank lending leading to higher interest rates affects manufacturing firms' demand for trade credit, measured as accounts payable relative to sales.[4]

[4] He uses aggregate data on manufacturing firms from the Quarterly Financial Reports (QFR) and annual firm-level data on 238 manufacturing firms from Compustat between 1975 and 1999.

Consistent with a trade credit channel, Nilsen finds that small public and private firms increase their payables in tight-money periods, as expected from firms with few financing alternatives. Additionally, he finds that large public firms without bond ratings increase their payables when money is tight whereas firms with bond ratings do not.

Choi and Kim (2005) consider how tighter monetary policy affects supply and demand for trade credit at S&P 500 firms and a sample of non-S&P 500 firms. They estimate firm-level supply and demand models on quarterly Compustat data.[5] Consistent with a trade credit channel, Choi and Kim find that tight money leads firms of both types to expand their supply of trade credit but that non-S&P 500 firms expand supply significantly more. This result differs from the Meltzer (1960) conception of the channel since non-S&P 500 firms have generally poorer credit market access than S&P 500 firms. Also at odds with Meltzer is their finding that tight money leads both firm types to expand their demand for trade credit by the same amount, rather than to a greater expansion at non-S&P 500 firms. Choi and Kim conclude that a trade credit channel exists but that the flow of trade credit is not unidirectional from firms with better credit market access to firms with poorer access.

Our paper also relates to work by Garcia-Appendini and Montoriol-Garriga (2013) and Yang (2011), who examine supply and demand for trade credit around the 2008 financial crisis. Garcia-Appendini and Montoriol-Garriga (2013) investigate supplier-client pairs using quarterly

[5] They measure trade credit demand and supply as payables-to-assets and receivables-to-assets, respectively, over a 1974-1997 sample period.

Compustat data.[6] They find that the crisis reduced the overall supply of trade credit but expanded the supply from suppliers with greater liquidity. They also find that the crisis generally increased the demand for trade credit especially at firms with poor credit market access as judged by several metrics including the Whited and Wu (2006) index. Yang (2011) studies supply and demand for trade credit at manufacturing firms using quarterly Compustat data between 2005 and 2009. He finds that once the crisis broke all firms expanded their supply and demand for trade credit.

Our paper complements work by Demiroglu et al. (2012), who study how changes in bank lending standards following monetary policy changes affect bank credit lines. Their sample comprises 2,141 private firms and a like number of public firms with similar attributes observed from 1993 to 2003. The sample firms are quite small, having median total assets of around $22m in year 2000 prices. Demiroglu et al. find that tighter lending standards affect public and private firms having established credit lines similarly however private firms lacking lines get significantly fewer new lines and increase accounts payable significantly more than public firms lacking credit lines.

Our paper advances the literature on trade credit and the transmission of US monetary policy in two ways. The first is through our empirical strategy: we test for the existence of a trade credit channel using firm-level quarterly data on essentially the universe of publicly-traded non-financial corporations. Use of firm-level data allows us to estimate firm-level models of trade credit supply and demand which control for other motives for using trade credit beyond the

[6] Their sample represents less than a quarter of firms with quarterly Compustat data. They measure the supply of trade credit as accounts receivable over sales and the demand for trade credit as accounts payable over cost-of-goods sold.

purely financial motive.[7] Use of quarterly data allows us to measure more precisely the link between change in monetary policy and change in the supply and demand for trade credit. Use of virtually all public non-financial firms yields a more comprehensive view of the total effect of monetary policy on trade credit flows.[8]

Our second contribution is our novel means of representing the credit market access of our sample firms, a key issue since the existence of a trade credit channel presupposes unequal access. We use two credit market access measures: the Z-Score index of Altman (1968) as modified by MacKie-Mason (1990) and the Whited and Wu (1996) index. Z-Score, which gauges the probability of bankruptcy in the coming 24 months, enjoys widespread use.[9] Molina and Preve (2009, 2012) show that financial distress changes firms' supply and demand for trade credit while decreasing firms' access to bank credit.[10] The Whited and Wu (2006) index measures difficulty in financing investment. Researchers use it to gauge constraints on firms'

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[7] See Mateut (2005) for a more detailed review of theories of trade credit.

[8] Our empirical approach stands in contrast to prior approaches. Gertler and Gilchrist (1993) and Oliner and Rudebusch (1996) use aggregate quarterly data on manufacturing firms. Nilsen (2002) uses both aggregate quarterly data and firm-level annual data on manufacturing firms. Choi and Kim (2005) use firm-level quarterly data on subsets of large and small firms from the Compustat universe. Demiroglu et al. (2012) use firm-level annual data on samples of private and public firms. Garcia-Appendini and Montoriol-Garriga (2013) and Yang (2011) use firm-level quarterly data on samples of manufacturing firms.

[9] As one illustration, Aktas et al. (2012) note that financial institutions use Z-score to judge firms' financial health in making lending decisions, supporting our use of Z-score as a measure of credit market access.

[10] Molina and Preve (2009, 2012) study trade credit at financially distressed firms from 1978-2000 using annual Compustat data. Increasing financial distress reduces firms' supply of trade credit (Molina and Preve, 2009) and increases their demand (Molina and Preve, 2012).

financial market access (e.g., Livdan et al., 2009).[11] We use Z-Score and WW-Score both as independent variables in our trade credit supply and demand models and as a means of stratifying our sample to produce subsamples of firms having similar credit market access.

The rest of the paper is organized as follows. Section II describes our estimation strategy. Section III presents the data and descriptive statistics. Section IV reports and discusses our empirical results. Section V concludes the paper.

II. Test Strategy

If a trade credit channel exists, a monetary policy tightening should expand both the supply of trade credit from financially strong firms with good credit market access and the demand for trade credit by financially weak firms with poor credit market access. We test this conjecture.

The conjecture implies the components of a test strategy: firm-level data spanning a time interval over which the stance of monetary policy changes; a measure of policy stance; measures of firms' credit market access; and econometric models of the supply and demand for trade credit. We employ these components as follows. We start with quarterly firm-level data over an extended time interval and define a zero-one indicator variable to identify tight money periods, D_TMP. We define two measures of a firm's credit market access, CMA1 and CMA2, and compute these measures for every firm each quarter. We sort the firms by CMA1 each quarter

[11] Most prior research on the trade credit channel uses firm size to measure credit market access (e.g., Gertler and Gilchrist, 1993; Oliner and Rudebusch, 1996; and Choi and Kim, 2005). Nilsen uses firm size, public or private status and presence of a bond rating. Demiroglu et al. (2012) uses public or private status. Yang (2011) uses presence of a bond rating. Garcia-Appendini and Montoriol-Garriga (2013) use presence of a bond rating along with six other measures, among them WW-score.

and define a set of indicators for the CMA1 quartile, D_CMA1_n, n=1,2,3,4, 4 indicating the poorest access. We repeat the process using CMA2. We then use CMA1 to stratify the whole sample into subsamples with different degrees of credit market access and use CMA2 to identify firms having different degrees of credit market access within these subsamples. Specifically, we designate the panels for which D_CMA1_1 =1 and D_CMA1_4 =1 as the unconstrained and constrained panels, respectively. Within each panel we again sort firms by CMA2 each quarter and define the quartile indicators D_CMA2_n, n=1,2,3,4, 4 indicating the poorest access. Finally we estimate reduced-form models of trade credit supply and trade credit demand, represented by the ratio of a firm's accounts receivable to assets, AR/A, and accounts payable to assets, AP/A, on data for the whole sample and for the unconstrained and constrained panels.

$$(1)\qquad \left(\frac{AR}{A}\right)_{it} = \alpha^s + \sum_{n=1}^{3} \beta_n^s\, D\_CMA2_{n,i,t-1} + \varphi^s\, D\_TMP_{t-1} + \sum_{n=1}^{3} \gamma_n^s\, D\_CMA2_{n,i,t-1} * D\_TMP_{t-1}$$
$$+ \sum_{k}^{K} \delta_k^s\, X_{k,i,t-1} + \sum_{n=1}^{3} \theta_n^s\, D\_Q_n + \mu_i^s + \varepsilon_{it}^s$$

$$(2)\qquad \left(\frac{AP}{A}\right)_{it} = \alpha^d + \sum_{n=1}^{3} \beta_n^d\, D\_CMA2_{n,i,t-1} + \varphi^d D\_TMP_{t-1} + \sum_{n=1}^{3} \gamma_n^d\, D\_CMA2_{n,i,t-1} * D\_TMP_{t-1}$$
$$+ \sum_{j}^{J} \delta_j^d\, W_{j,i,t-1} + \sum_{n=1}^{3} \theta_n^d\, D\_Q_n + \mu_i^d + \varepsilon_{it}^d$$

Subscripts i and t denote firm i and quarter t. Superscripts s and d denote parameters in the trade credit supply and demand equations, respectively. $X_{k,i,t-1}$, k = 1,.., K and $W_{j,i,t-1}$, j = 1,..,J are firm-specific determinants of firm i's trade credit supply and demand, respectively. We observe these determinants one quarter earlier than the dependent variable, at time t-1, to avoid endogeneity bias. D_Q_n, n = 1,2,3,4, are binary indicators of calendar quarters to control for seasonality effects; D_Q_4 is the omitted quarter. Similarly D_CMA2_4, the indicator for CMA2 quartile 4, having the poorest credit market access, is the omitted category. μ_i^s and μ_i^d are unobservable, time-invariant fixed firm effects in the supply and demand equations, respectively. ε_i^s and ε_i^d are error terms assumed to have all desirable properties.

Initial evidence on the trade credit channel comes from the estimated coefficients of the credit market access variables D_CMA2_n, n=1,2,3, namely β_n^s and β_n^d, n=1,2,3. If a channel exists, supplier firms with good credit market access offer trade credit to support purchases by buyer firms with poorer access. Offers of trade credit should decline as supplier firms' credit market access declines while demand for trade credit should rise as buyer firms' credit market access declines. Estimates of (1), the trade credit supply equation, will support the existence of a trade credit channel if the estimated coefficients of D_CMA2_n decline in n, i.e., $\beta_1^s > \beta_2^s > \beta_3^s > 0$, and are larger in the unconstrained panel. Estimates of (2), the trade credit demand equation, will support the existence of a trade credit channel if the estimated coefficients rise in n, i.e., $\beta_1^d < \beta_2^d < \beta_3^d < 0$, and have larger absolute values in the constrained panel.[12]

Further evidence on the trade credit channel comes from the estimated coefficients of the tight monetary policy indicator D_TMP_{t-1}, φ^s and φ^d. If a channel exists tighter policy expands the supply of trade credit, especially at firms with the best credit market access, and expands the demand for trade credit particularly at firms with the poorest credit market access. Thus when we estimate (1) and (2) on separate panels of unconstrained and constrained firms, the estimated φ^s

[12] Our choice of D_CMA2_4 as the omitted category affects the interpretation of β_n^s and β_n^d, n=1,2,3. β_n^s (β_n^d) is the effect on a firm's trade credit supply (demand) of being in credit market access category n, n=1,2,3, relative to being in category 4. Firms in category 4 should supply the least trade credit (hence we expect $\beta_n^s > 0$) and demand the most trade credit (hence we expect $\beta_n^d < 0$).

should be larger in the former panel and the estimated φ^d should be larger in the latter. When we estimate (1) and (2) on data for the whole sample the estimates of φ^s and φ^d should be positive.[13]

Final evidence on the trade credit channel comes from the estimated coefficients of the tight money/credit market access interaction terms $\mathrm{D\_CMA2}_{n,i,t-1} * \mathrm{D\_TMP}_{t-1}$, n = 1,2,3, namely γ_n^s and γ_n^d, n=1,2,3. If a trade credit channel exists, tighter policy expands the supply and demand for trade credit with supply increasing in firms' credit market access and demand decreasing in firms' credit market access. Estimates of (1) will support the existence of a channel if $\gamma_1^s > \gamma_2^s > \gamma_3^s > 0$ and if the γ_n^ss are larger in the unconstrained panel. Estimates of (2) will support the channel and if $\gamma_1^d < \gamma_2^d < \gamma_3^d < 0$ and if the γ_n^ds have larger absolute values in the constrained panel.[14]

[13] In our modeling, φ^s and φ^d measure both the effect of tight money on supply and demand for trade credit common to all firms and the effect of tight policy on firms in the omitted credit market access category, $\mathrm{D\_CMA2}_4$ =1. In general, φ^s and φ^d could be negative if tight policy dampens economic activity. Garcia-Appendini and Montoriol-Garriga (2013) find that the overall supply of trade credit fell at the onset of the financial crisis, implying $\varphi^s < 0$ for their sample.

[14] Our modeling approach imparts the following interpretations on the coefficients. β_n^s (β_n^d) shows the effect on trade credit supplied (demanded) of a firm being in credit market access quartile 1, 2, or 3 relative to quartile 4 in a loose-money period (i.e. $\mathrm{D\_TMP}_{-1}$=0); φ^s (φ^d) shows the common effect on trade credit supplied (demanded) of a tight-money period; and $\beta_n^s + \gamma_n^s$ ($\beta_n^d + \gamma_n^d$) shows the effect on trade credit supplied (demanded) of a firm being in credit market access quartile 1, 2, or 3 relative to quartile 4 in a tight-money period.

III. Data

A. Sample Design and Description of Variables

We start with the universe of firms in the quarterly Compustat database from 1988 through 2008.[15] We exclude financial service firms (SIC Code 60-69), utility firms (SIC Code 49), and non-classified firms (SIC Code 99). We also omit firms with negative or zero assets and quarterly sales of less than $5 million. The panel is unbalanced which allows for entry and exit, partially mitigating concerns about sample selection and survival bias. The whole sample comprises 303,633 firm quarters.

Following Bernanke and Blinder (1992) we use the effective federal funds rate to distinguish quarters of tight and loose monetary policy. Specifically, we define a tight-money quarter as one with an end-of-quarter effective fed funds rate of 5% or more. We define a tight monetary policy indicator, D_TMP, to be one in tight-money quarters and zero otherwise. By this definition 42 of the 84 quarters from 1988 to 2008 are tight-money quarters.

We use the Whited and Wu (2006) index as our first credit market access measure, CMA1. WW-score measures a firm's difficulty in financing investment, with larger values indicating greater difficulty and, hence, poorer credit market access. Table 1 reports the definition of WW-score. We sort our sample firms by CMA1 each quarter and identify the least and most constrained quartiles, which we call the unconstrained and constrained panels, respectively.[16] We permit firms to move in and out of these panels over time as their financial performances

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[15] We start our sample period in 1988 because this is the first year quarterly Compustat data are complete for most firms. We include both active and inactive firms.

[16] For all firms in the unconstrained (constrained) panel $D\_CMA1_1 = 1$ ($D\_CMA1_4 = 1$).

warrant. Our unconstrained (constrained) panels comprise 68,482 firm-quarters (70,989 firm-quarters).

**** Table 1 about here ****

We use Altman's (1968) Z-score as modified by MacKie-Mason (1990) as our second credit market access measure, CMA2. Z-score indexes the likelihood that a firm declares bankruptcy in the coming 24 months, with larger values implying more probable bankruptcy and, hence, poorer credit market access.[17] Table 1 reports the definition of Z-score. We sort our whole sample and our unconstrained and constrained panels by Z-score each quarter and define Z-score quartile indicators, D_CMA2_n , n = 1,2,3,4. We allow firms to migrate across quartiles as their financial conditions dictate. D_CMA2_1 =1 (D_CMA2_4 =1) indicates a firm in the best (poorest) credit market access quartile

We represent firm-specific determinants of trade credit supply using eight variables, the X_ks in (1): LnA_{-1}, $LnAge_{-1}$, Tobin's Q_{-1}, $Sales/A_{-1}$, CS/A_{-1}, and AP/A_{-1}, $D\_TMP*Sales/A_{-1}$ and D_TMP*CS/A_{-1}. Assets (LnA_{-1}) and firm age ($LnAge_{-1}$) characterize a firm's reputation. Firms with established reputations need not extend trade credit to attract sales unlike relatively unknown firms (Deloof and Jegers, 1996; Long et al, 1993; Murfin and Njoroge, 2015), hence we expect negatively signed coefficients on LnA_{-1} and $LnAge_{-1}$. Tobin's Q_{-1} represents growth opportunities (Deloof and Jegers, 1999). Firms with greater opportunities may try to raise the cash to exploit them by selling their products on credit, leading to a positive coefficient on

[17] We reverse the algebraic sign of Z-score so that higher values imply more probable bankruptcy and, hence, poorer credit market access, similar to WW-score.

Tobin's Q_{-1}. Accounts receivable will be tied to prior-quarter sales when firms routinely extend some trade credit to facilitate transactions and when quarterly sales show positive time trends, producing a positive coefficient on $Sales/A_{-1}$. Production lead times create correlations between current-quarter accounts receivable and prior-quarter balance sheet accounts. Specifically, firms with greater accounts receivable will have previously reduced cash and increased accounts payable to buy inputs to production, producing a negative coefficient on CS/A_{-1} and a positive coefficient on AP/A_{-1}. We include interactions between the monetary policy variable D_TMP_{-1} and the control variables $Sales/A_{-1}$ and CS/A_{-1} to permit them to exert a different effect on accounts receivable in tight money periods.

We represent firm-specific determinants of trade credit demand using twelve variables, the W_js in (2): LnA_{-1}, $LnAge_{-1}$, Tobin's Q_{-1}, $Sales/A_{-1}$, CS/A_{-1}, INV/A_{-1}, AR/A_{-1}, CF/A_{-1}, LTD/A_{-1}, $D\_TMP*Sales/A_{-1}$, $DTMP*CS/A_{-1}$, and D_TMP*CF/A_{-1}. Assets and firm age again characterize a firm's reputation. Firms with established reputations demand less trade credit as they have access to better funding sources, producing negative coefficients on LnA_{-1} and $LnAge_{-1}$. Tobin's Q_{-1} again represents growth opportunities. Firms with more opportunities are more likely cash constrained and thus in more need of trade credit, suggesting a positive coefficient for Tobin's Q_{-1}. Positive correlation between current and prior-quarter sales should lead greater prior-quarter sales to raise current-quarter demand for trade credit to finance purchases of product inputs, implying a positive coefficient on $Sales/A_{-1}$. For firms which match the maturities of short-term assets and liabilities, higher levels of prior-quarter cash (CS/A_{-1}), inventories (INV/A_{-1}), and accounts receivable (AR/A_{-1}) should produce greater demand for trade credit, leading to positive coefficients on CS/A_{-1}, INV/A_{-1}, and AR/A_{-1}. Greater cash flow relative to assets (CF/A_{-1}), defined as net income plus depreciation divided by assets, facilitates purchases of product inputs and reduces firms' need for trade credit, leading us to expect a negatively signed coefficient on CF/A_{-1}. Better access to long-term credit, evidenced by a larger long-term-debt-to-asset ratio (LTD/A_{-1}), reduces the need for trade credit, producing a negative coefficient on LTD/A_{-1}. Interactions between the monetary policy variable D_TMP_{-1} and the controls $Sales/A_{-1}$, CS/A_{-1}, and CF/A_{-1}, allow these controls to impact accounts payable differently when money is tight.

B. Descriptive Statistics

Table 2 presents descriptive statistics for the whole sample and for the unconstrained and constrained panels. Unconstrained firms are significantly larger, faster growing and more established, on average, than constrained firms. In the whole sample mean book-value assets and sales are \$2.7b and \$0.6b, respectively, in 2008 dollars; in the unconstrained and constrained panels mean assets and sales are, respectively, more than double the whole sample means and less than half the whole sample means. In the whole sample annual sales growth (SG) averaged nearly 2% over the 21-year sample period; in the unconstrained and constrained panels annual sales growth averaged 4.6% and -0.1%, respectively. In the whole sample mean firm age is 11.4 years; in the unconstrained and constrained panels mean age is 13.3 years and 11 years, respectively.

**** Table 2 about here ****

Descriptive statistics for CMA1 and CMA2 (WW-score and Z-score, respectively) confirm that firms in the unconstrained panel have significantly better credit market access than firms in the constrained panel. In the unconstrained panel mean CMA1 is -0.83 with a standard deviation of 0.28; the same statistics for the constrained panel are -0.26 and 0.31, respectively. Similarly, mean CMA2 is -0.71 with a standard deviation of 0.98 in the unconstrained panel, and -0.31 and 1.64, respectively, in the constrained panel.

Mean AR/A and AP/A are significantly different in the whole sample (0.18 and 0.10, respectively), a counter-intuitive result since it implies that the sample firms extended more trade credit than they received, a clear violation of accounting identities since extending trade credit creates an account receivable for the creditor firm and an identical-size account payable for the debtor firm. We contend that the wedge between mean AR/A and mean AP/A mainly reflects the omission of private firms from the Compustat universe. Specifically, private firms have limited sources of funds, making them more likely than public firms to demand trade credit and less likely to supply it. Consequently the whole-sample mean AR/A represents trade credit supplied by the average public firm to both public and private firms whereas the whole-sample mean AP/A represents trade credit demanded by the average public firm primarily from other (mainly public) firms.

Mean AR/As are also significantly different in the constrained and unconstrained panels (0.20 and 0.16, respectively), also a counter-intuitive result as it suggests that constrained firms supply more trade credit. We maintain that this difference is due in part to differences in the composition of assets at constrained and unconstrained firms. In particular, unconstrained firms can use their superior credit market access to obtain more readily long-term loans with which to buy such long-term assets as plant and equipment or to engage in mergers and acquisitions. Two pieces of evidence from Table 2 support this reasoning. First, the mean long-term debt-to-asset ratio (LTD/A_{-1}) is significantly larger in the unconstrained panel (23% vs. 19%). Second, the mean cash-to-asset ratio (CS/A_{-1}) and mean inventory-to-asset ratio (INV/A_{-1}) are smaller in the unconstrained panel (11% and 16% vs. 15% and 19%).

**** Table 3 about here ****

Table 3 shows the effect of credit market access on AR/A and AP/A. The first column shows CMA1 deciles arranged in order of increasing difficulty of financing investment and accessing credit markets; the second column shows mean CMA1 in each decile. Mean and median AR/A rise with declining credit market access, reflecting in part the difficulty financially constrained firms face in accessing credit markets to finance long-term assets. Mean and median AP/A also rise with declining credit market access as expected when poorer credit market access forces greater reliance on trade credit, although the rise is much less pronounced than the rise in mean and median AR/A.

**** Table 4 about here ****

Table 4 reports mean AR/A, AP/A, CMA1 and CMA2 by year for the whole sample and for the unconstrained and constrained panels. Mean AR/A declines over the sample period in all three samples. Mean AP/A declines in the whole and the constrained samples but shows little movement in the unconstrained sample. In every year mean AR/A for the constrained panel equals or exceeds mean AR/A for the unconstrained panel. An analogous statement applies to AP/A, except in 2008. Mean CMA1 fluctuates over the sample period, reflecting changes in the difficulty of financing investment and accessing capital credit market over the business cycle. In the whole sample mean CMA1 falls to -0.70 in 1991 and to -0.72 in 2001 near the cyclical peaks in July 1990 and in March 2001. Mean CMA2 rises over the sample period in all three samples, implying a general rise in financial distress and a general decline in credit market access. Mean CMA2 is consistently lower in the unconstrained sample than the constrained sample, however.

**** Table 5 about here ****

Table 5 presents statistics on CMA1, AR/A, and AP/A stratified by the stance of monetary policy. Panel A shows for the whole sample that the average firm is significantly more constrained when policy is tight: mean CMA1 in loose- and tight-money periods are -0.57 and -0.53, respectively. Panel B shows that firms use significantly more trade credit when policy is tight: in tight-money periods mean AR/A and mean AP/A are larger in the whole sample and in the unconstrained and constrained panels.

IV. Empirical Results and Discussion

We begin by addressing the problem of modeling differences among firms over time. Hausman tests for random versus fixed-effects specifications yield large values for all the models, leading us to reject the random-effects specification. All model estimates presented below include unreported fixed-effects constants. The fixed-effects specification prevents us from including industry indicators among our explanatory variables. We explore the possibility of industry effects as a robustness check in Section IV.C.

A. Supply of Trade Credit

Table 6 reports estimates of equation (1) and robust standard errors to control for heteroskedasticity and within-cluster (-firm) correlation.[18] Estimates of (1) produced by the whole sample (WS) and by the unconstrained (U) and constrained (C) subsamples appear in columns (1)-(3), respectively.

**** Table 6 about here ****

The whole sample model estimate strongly supports the existence of a trade credit channel (column 1). In loose-money periods firms supply trade credit commensurate with their credit market access, as shown by the positive and decreasing estimated coefficients of D_CMA2_n, n = 1,2,3. More specifically, firms with the best access (i.e., D_CMA2_1 =1) supply nearly 12% more trade credit than firms with the poorest access (i.e., D_CMA2_4 =1, the excluded category) and firms with the next-to-best and next-to-poorest access (i.e., D_CMA2_2 =1 and D_CMA2_3 =1, respectively) supply 9% more and 4.5% more, respectively.[19] Tight policy leads firms to expand the supply of trade credit: the positive estimated coefficient of D_TMP_{-1} implies that firms

[18] Stock and Watson (2008) note that while the standard heteroskedasticity-robust estimator is inconsistent for the fixed effects estimator, the cluster-robust estimator is consistent.

[19] Table 5 Panel B shows a mean AR/A of 0.179 for the whole sample. In Table 6 column (1) the estimated coefficient of D_CMA2_1 is 0.021. Hence compared with firms having the poorest credit market access, firms with the best credit market access supply (100 x 0.021/0.179 =) 11.7% more trade credit.

increase their accounts receivable by about 3.4% on average.[20] In addition, tight policy leads firms to expand the supply of trade credit by amounts commensurate with their credit market access, as indicated by the positive and generally decreasing estimated coefficients of $D\_TMP_{-1}*D\_CMA2_n$, n = 1,2,3. The estimated coefficients imply that, relative to firms with the poorest credit market access, firms with the best, next-to-best and next-to-poorest credit market access expand their accounts receivable by additional amounts averaging 2.8%, 2.8%, and 1.7%, respectively.

The estimated coefficients of the remaining independent variables conform to expectations. Smaller and younger firms extend more trade credit per dollar of assets (LnA_{-1} and $LnAge_{-1}$ have significant negative coefficients) consistent with a strategy of offering trade credit to compensate for unknown reputations. Firms with greater growth opportunities offer more trade credit (Tobin's Q_{-1} has a significant positive coefficient) consistent with promoting sales to alleviate cash constraints. Firms with greater prior-quarter sales supply more trade credit in the current quarter ($Sales/A_{-1}$ has a significant positive coefficient) as expected when firms offer some trade credit to facilitate transactions and when sales show positive time trends. These firms cut back their supply of trade credit when monetary policy is tight, however ($D\_TMP_{-1}*Sales/A_{-1}$ has a significant negative coefficient). Firms with less prior-quarter cash supply more trade credit (CS/A_{-1} has a significant negative coefficient) consistent with depleting cash to purchase product inputs, produce product and support credit sales. Tighter policy has no measurable impact on this process ($D\_TMP_{-1}*CS/A_{-1}$ has a statistically insignificant coefficient). Greater prior-quarter

[20] Table 5 Panel B shows a mean AR/A of 0.179 for the whole sample. In Table 6 column (1) the estimated coefficient of D_TMP is 0.006. Hence in tight money periods, the sample firms expand their accounts receivable by about (100 x 0.006/0.179 =) 3.4%.

accounts payable support greater purchases of product inputs leading to greater current-quarter credit sales (AP/A_{-1} has a significant positive coefficient).

Further support for a trade credit channel comes from the trade credit supply models estimated on the unconstrained and constrained panels (columns 2 and 3, respectively). In loose-money periods firms with the best credit market access supply the most trade credit and supply decreases with declining market access: in columns (2) and (3) the coefficients of D_CMA2_n, n = 1,2,3, are significant, positive and declining in n. More precisely, relative to firms with the poorest credit market access firms with the best access (i.e., D_CMA2_1 =1) supply 14.6% (10.5%) more trade credit in the unconstrained (constrained) subsample.[21] Analogously, firms with the next-to-best credit market access (i.e., D_CMA2_2=1) supply 10% (10.3%) more trade credit than firms with the poorest access in the unconstrained (constrained) subsample, while firms with the next-to-poorest credit market access (i.e., D_CMA2_3=1) supply 4% (5.5%) more trade credit in the unconstrained (constrained) subsample. In column (2) the estimated coefficients of the D_CMA2_n terms do not consistently exceed the estimated coefficients in column (3), contrary to expectations. Consistent with expectations, in tight-money periods unconstrained firms supply more trade credit but constrained firms do not: the estimated coefficient of D_TMP_{-1} is positive and significant in column (2) but statistically insignificant in column (3). The estimated coefficient of D_TMP_{-1} in column (2) implies that tight money leads

[21] Table 5 Panel B shows a mean AR/A of 0.151 for unconstrained firms in loose-money periods. In Table 6 column (2) the estimated coefficient of D_CMA2_1 is 0.022. Hence compared with unconstrained firms having the poorest credit market access, unconstrained firms with the best credit market access supply (100 x 0.022/0.151 =) 14.6% more trade credit.

unconstrained firms to expand their accounts receivable by an average of 2.5%.[22] Also consistent with expectations, tight-money periods lead unconstrained firms – but not constrained firms – to supply more trade credit in proportion to their credit market access: the estimated coefficients of $D\_TMP_{-1}*D\_CMA2_n$, n=1,2,3 are significant, positive and declining in n in column (2) but are statistically insignificant in column (3). The estimated coefficients in column (2) imply that tight money leads unconstrained firms with the best, next-to-best and next-to-poorest credit market access to add about 5.5%, 5.0% and 3.7%, respectively, to their usual supplies of trade credit.[23]

The estimated coefficients of the remaining explanatory variables are qualitatively similar to the estimates in column (1) with only slight differences. A one-percent increase in prior-quarter sales to assets leads to a 23.8% increase in accounts receivable to assets at constrained firms versus a 16.5% increase at unconstrained firms.[24] The larger elasticity may mirror constrained firms' greater need to use trade credit to facilitate transactions. This need does not abate in tight-money times: the coefficient of $D\_TMP_{-1}*Sales/A_{-1}$ is statistically zero for constrained firms instead of negative as for unconstrained firms. Constrained firms also recycle less trade credit: a

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[22]In Table 5 Panel B the mean AR/A for unconstrained firms is 0.157. In Table 6 column (2) the coefficient of D_TMP is 0.004. Hence in tight money times, unconstrained firms expand accounts receivable by (100 x 0.004/0.157 =) 2.5% on average.

[23] Table 5, Panel B shows a mean AR/A of 0.163 for unconstrained firms in tight money periods. In Table 6 column (2) the coefficient of $D\_TMP*D\_CMA2_1$ is 0.009. Thus when money becomes tight, the least constrained firms add (100 x 0.009/0.163=) 5.5% to their accounts receivable.

[24] Table 2 shows that constrained firms have a mean AR/A and mean $Sales/A_{-1}$ of 0.20 and 0.34, respectively. In Table 6 column (3) the coefficient of $Sales/A_{-1}$ is 0.14. Thus the elasticity of accounts receivable to assets with respect to sales to assets is (100 x 0.14 x 0.34/ 0.20 =) 23.8%

one-percent increase in prior-quarter accounts payable to assets raises accounts receivable to assets by about 10% at constrained firms versus nearly 15% at unconstrained firms.

B. Demand for Trade Credit

Table 7 reports estimates of equation (2) produced by the whole sample and by the unconstrained and constrained panels. The estimates appear in columns (1)-(3) respectively. [25]

**** Table 7 about here ****

The model estimate for the whole sample (column 1) shows weak evidence of a trade credit channel. Consistent with a channel, in loose-money periods firms' demand for trade credit is decreasing in credit market access, as shown by the negative and increasing estimated coefficients of D_CMA2_n, n=1,2,3. In particular, firms with the best credit market access (i.e., firms with D_CMA2_1 =1) demand about 27% less trade credit than firms with the poorest access (i.e., firms with D_CMA2_4 =1, the excluded category) while firms with the next-to-best and next-to-poorest access (i.e., firms with D_CMA2_2 =1 and D_CMA2_3=1, respectively) demand 17.5% less and about 10% less, respectively.[26] Inconsistent with a channel, tight monetary policy has

[25] In our models β_n^d, n=1,2,3, measures trade credit demand by firms in credit market access category n relative to category 4 in loose-money periods, φ^d measures the change in trade credit demand in tight-money periods common to all firms including firms in category 4, and γ_n^d, n=1,2,3, measures the change in trade credit demand in tight-money periods by firms in category n relative to firms in category 4.

[26] Table 5 Panel B reports the mean AP/A for the whole sample as 0.103. In Table 7 column (1) the coefficient D_CMA2_1 is -0.028. Hence relative to firms having the poorest credit market access, firms with the best access demand (100 x -0.028/0.103 =) 27.2% less trade credit.

no measurable effect on the common demand for trade credit by the sample firms: the estimated coefficient of D_TMP_{-1} is positively signed but statistically indistinguishable from zero. The interactive effect of tight monetary policy and credit market access is consistent with a trade credit channel but this effect is small. Specifically, in tight-money periods the marginal demand for trade credit is decreasing in firms' credit market access, as shown by the negative and increasing estimated coefficients of $D\_TMP_{-1}*D\_CMA2_n$, n=1,2,3, but only the marginal demand by firms with the best access differs statistically from the demand by firms with the poorest access; this difference is less than 3%.[27]

Most of the other independent variables have estimated coefficients which conform to expectations. Larger firms demand less trade credit (the coefficient of LnA_{-1} is negative and significant), consistent with established reputations affording better credit market access and less reliance on trade credit. Firms with greater growth opportunities demand more trade credit (the coefficient of Tobin's Q_{-1} is positive and significant) as they are typically cash constrained. Firms with greater prior-period sales have greater current-period demand for trade credit with which to buy production inputs (the coefficient of $Sales/A_{-1}$ is positive and significant). Tight money has no measurable effect on this demand (the coefficient of $D\_TMP_{-1}*Sales/A_{-1}$ is statistically insignificant). As expected of firms which match the maturities of their assets and liabilities, firms with more prior-period short-term assets – cash, accounts receivable and inventories – demand more trade credit in the current period (the coefficients of CS/A_{-1}, AR/A_{-1} and INV/A_{-1} are positive and significant). Tight money intensifies the relationship between

[27]Table 5 Panel B reports the mean AP/A for the tight money periods as 0.106. In Table 7 column (1) the coefficient $D\_TMP*D\_CMA2_1$ is -0.003. Hence relative to firms having the poorest credit market access, firms with the best access reduce their demand for trade credit by (100 x -0.003/0.106 =) 2.8%.

prior-period cash and current-period trade credit demand (the coefficient of $D\_TMP_{-1}*CS/A_{-1}$ is positive and significant). Firms with greater cash flow demand less trade credit especially in tight-money times (the coefficients of CF/A_{-1} and $D\_TMP_{-1}*CF/A_{-1}$ are negative and significant) as internal funds cost less than trade credit. Similarly, firms with greater access to long-term debt markets demand less trade credit (the coefficient of LTD/A_{-1} is negative and significant). Only the coefficient of $LnAge_{-1}$ defies expectations: older firms with reputations which should give them better credit market access demand more trade credit instead of less (the coefficient of $LnAge_{-1}$ is positive and significant).

Estimated trade credit demand models for the unconstrained and constrained firms (columns 2 and 3, respectively) also show weak evidence of a trade credit channel. Consistent with a channel, the model estimates for both panels show that firms' demand for trade credit is decreasing in credit market access during loose-money periods. Also consistent with a channel, the estimated coefficients of the credit market access variables have smaller absolute values in the unconstrained panel. Specifically in the unconstrained panel, firms with the best credit market access, next-to-best access and next-to-poorest access demand 20% less, 11% less and 6% less trade credit than firms with the poorest market access, respectively. [28] The analogous metrics in the constrained panel are 32%, 22% and 13%. Tight monetary policy has no measurable effect on the common demand for trade credit by firms in either the unconstrained or constrained panels: the estimated coefficients of D_TMP_{-1} are statistically insignificant instead of being positive and larger in column (3) as predicted. The estimated coefficients of the tight

[28] Table 5 Panel B shows a mean AP/A of 0.097 for unconstrained firms. In Table 7 column (2) the estimated coefficient of D_CMA2_1 is -0.019. Hence compared with unconstrained firms having the poorest credit market access unconstrained firms with the best access demand (100 x -0.019/0.097) 19.6% less trade credit.

money/credit market access interaction terms also contravene expectations. In particular in the unconstrained panel, tight policy reduces by about 4% trade credit demand by firms with the best credit market access relative to firms with the poorest access while leaving demand by firms in the two other access categories unchanged. In the constrained panel all of the tight money / credit market access interaction terms have statistically insignificant coefficients instead of having coefficients larger in absolute value than in the unconstrained panel.

The other explanatory variables in the unconstrained and constrained model estimates have coefficients similar to those for the whole sample estimate, with three exceptions. First, the prior-quarter cash-to-assets ratio (CS/A_{-1}) has coefficients with opposite algebraic signs: positive in the unconstrained model estimate as expected when firms match the maturities of their assets and liabilities but negative in the constrained model estimate. The latter coefficient, which implies that a one-percent cash ratio increase reduces trade credit demand by 1.5%, suggests that constrained firms treat cash and trade credit as substitute methods of purchasing production inputs, as might be expected when firms are cash constrained and trade credit is expensive. [29] Tight-money periods weaken this substitutability, as might be expected when firms are especially cash-constrained: the positive coefficient of $D\_TMP_{-1}*CS/A_{-1}$ in column (3) moves the elasticity of trade credit demand with respect to the cash ratio from -1.5% to -0.4%, near zero.[30] Second, the prior-quarter long-term debt-to-asset ratio has coefficients with differing

[29] Table 2 shows that constrained firms have mean AP/A_{-1} and CS/A_{-1} of 0.11 and 0.15, respectively. In Table 7 column (3) the estimated coefficient of CS/A_{-1} is -0.011. Hence a one-percent increase in CS/A_{-1} decreases AP/A_{-1} by (100 x -0.011 x 0.15/0.11 =) 1.5% . For unconstrained firms the same increase in CS/A_{-1} raises AP/A_{-1} by 1.5%.

[30] In Table 7 column (3) the estimated coefficient of D_TMP* CS/A_{-1} is 0.008. Hence in tight-money periods a one-percent increase in CS/A_{-1} decreases AP/A_{-1} by (100 x (-0.011 + 0.008) x 0.15/0.11=)0.4%. For unconstrained firms the same increase in CS/A_{-1} raises AP/A_{-1} by more than 2%.

statistical significance. Specifically the long-term debt ratio has a negatively signed coefficient in the unconstrained model estimate, consistent with our prior belief that firms with good credit market access demand less trade credit, but a statistically insignificant coefficient in the constrained model estimate, as might be expected at firms having little long-term debt due to poor credit market access. Finally the tight money / cash flow interaction term, D_TMP_{-1}* CF/A_{-1}, has statistically insignificant coefficients in both the unconstrained and constrained model estimates despite having a significant negative coefficient in the whole sample estimate. Thus tighter money has no measurable effect on the substitutability between cash flow and trade credit as methods of buying production inputs.

C. Robustness Checks

We check the robustness of our main results by modifying and re-estimating our trade credit supply and demand models.[31]

**** Tables 8 and 9 about here ****

Tables 8 and 9 report estimated trade credit supply and demand models which replace the binary tight-money indicator, D_TMP_{-1} , with the effective federal funds rate, FFR_{-1}, a continuous variable.[32] The supply model estimates in Table 8 are similar to the original estimates in Table 6, with two exceptions. First, in loose-money periods the supply of trade

[31] All of the estimated models described in this section but not reported are available upon request.

[32] Characterizing monetary policy with FFR presumes that an incremental change in the federal funds rate has a constant impact on trade credit supplied or demanded irrespective of its level. In contrast, characterizing policy with D_TMP_{-1} presumes that the impact of policy depends only on whether the fed funds rate is below or above 5%.

credit from unconstrained firms is less sensitive to credit market access: in column (2) the coefficients of D_CMA2_n, n=1,2,3 are positive and declining in n but notably smaller than in Table 6. Further, these coefficients are smaller than the corresponding coefficients for constrained firms in column (3), contrary to expectations. Second, a monetary policy tightening expands the supply of trade credit by constrained firms: the coefficient of FFR_{-1} in column (3) is significant and positive, unlike the coefficient of D_TMP_{-1} in Table 6, column (3). Further, the coefficient of FFR_{-1} is larger for constrained firms than for unconstrained firms (0.003 versus 0.001), contrary to expectations. In Table 9 the demand models using FFR_{-1} as the tight-money indicator are very similar qualitatively to the original estimates in Table 7.

**** Tables 10 and 11 about here ****

Tables 10 and 11 report estimated trade credit supply and demand models which characterize credit market access with the continuous variable CMA2 (Z-score) in place of the credit market access quartile-indicators D_CMA2_n, n=1,2,3,4.[33] The supply model estimates in Table 10 strongly support the existence of a trade credit channel and are qualitatively similar to the original estimates in Table 6.[34] Specifically, in loose-money periods trade credit supply is

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[33] Representing credit market access with the continuous variable CMA2 presumes that an incremental change in credit market access has a constant impact on trade credit supplied or demanded. Representing credit market access using the quartile indicators D_CMA2_n, n=1,2,3 allows an incremental change in credit market access to have different impacts. The estimated coefficients of CMA2 and D_CMA2_n should have opposite algebraic signs due to our choice of CMA2 quartile 4 as the excluded category.

[34] Using CMA2 in place of the quartile-indicators D_CMA2_n, n=1,2,3, changes the conditions supporting a trade credit channel as follows: $\beta^s,\gamma^s<0$; $\beta^d,\gamma^d>0$; $|\beta^s|,|\gamma^s|$ larger for unconstrained firms; $|\beta^d|,|\gamma^d|$ larger for constrained firms.

increasing in credit market access particularly for unconstrained firms ($CMA2_{-1}$ has negative coefficients); a policy tightening increases trade credit supply from all firms but particularly from unconstrained firms (D_TMP_{-1} has positive coefficients); and marginal trade credit supply in tight-money periods is increasing in credit market access particularly for unconstrained firms ($D\_TMP_{-1}*CMA2_{-1}$ has negative coefficients). In Table 11 the demand model estimates are qualitatively similar to the original model estimates in Table 7 and less supportive of a trade credit channel. Specifically, like the original model estimates and consistent with a trade credit channel, trade credit demand is decreasing in credit market access in loose-money periods especially for constrained firms ($CMA2_{-1}$ has positive estimated coefficients), as is marginal demand in tight-money periods ($D\_TMP_{-1}*CMA2_{-1}$ has positive coefficients). Like the original model estimates but inconsistent with a trade credit channel, tight-money periods do not measurably increase the general demand for trade credit demand (D_TMP_{-1} has statistically insignificant coefficients) and marginal trade credit demand is larger from unconstrained firms ($D\_TMP_{-1}*CMA2_{-1}$ has a larger estimated coefficient in column 2 than in column 3).

In addition to the model estimates we report in Tables 6-11 which use a fixed-effects specification, we estimate analogous models which use a random-effects specification. This latter specification allows the inclusion of industry indicators in the trade credit supply and demand models, a potentially important modification. However, the estimated models with random-effects are qualitatively similar to the estimated models with fixed-effects.

Equations (1) and (2) use quarterly indicator variables to control for seasonality effects however firms in different industries may face different seasonal sales patterns and may choose to smooth them out using trade credit (Bougheas et al., 2009). We explore this possibility in two ways. First, we re-estimate fixed-effect models augmented by a set of interactions between quarterly indicators and industry indicators. [35] Second, we re-estimate random-effects models which include quarterly indicators, industry indicators, and interactions between them. These re-estimated models are all qualitatively similar to the estimates reported in Tables 6 and 7 except for exhibiting stronger evidence favorable to CMA1 and CMA2 as determinants of trade credit supply and demand.

D. Discussion

The estimated whole-sample trade credit supply and demand models, reported in column (1) of Tables 6 and 7 respectively, imply that a monetary policy tightening, represented by a change in D_TMP from 0 to 1, increases the net supply of trade credit from public firms, represented by AR/A_{-1} - AP/A_{-1}, and directs it from firms with better credit market access towards firms with poorer access. Specifically, a tightening increases the common supply of trade credit (the point estimate of φ^s is .006) without measurably affecting the common demand for trade credit (the point estimate of φ^d is zero) thus expanding the common net supply of trade credit ($\varphi^s - \varphi^d$ = .006 > 0). Additionally, a tightening both increases the supply of trade credit from firms having the best credit market access relative to firms having the poorest access (the point estimate of γ_1^s is .005) and decreases their demand (the point estimate of γ_1^d is -.003) thus expanding their net supply of trade credit ($\gamma_1^s - \gamma_1^d$ = .008 > 0). Analogously, a tightening expands the net supply of trade credit from firms having the next-to-best and next-to-poorest credit market access relative to firms having the poorest access ($\gamma_2^s - \gamma_2^d$ = .005 > 0 and $\gamma_3^s - \gamma_3^d$ = .003 > 0, respectively).

[35] Collinearity prevents the inclusion of industry dummies in a fixed-effects model, hence we control for industry effects by controlling unobservable individual effects at the firm level. Collinearity does not prevent inclusion of industry/quarterly interactions terms, however.

Thus our empirical evidence supports the existence of a trade credit channel as characterized by Meltzer (1960).

Our findings differ somewhat from those of Choi and Kim (2005) who also investigate the existence of a trade credit channel.[36] They use the 3-month Treasury bill rate, R, to represent the stance of monetary policy in trade credit supply and demand models estimated on quarterly Compustat data for S&P500 firms and for a similar number of other public firms.[37] In trade credit demand equations estimated for the two panels they find that the coefficients of R are significant, positive and approximately equal, suggesting that tight policy increases about equally trade credit demand by S&P500 and non-S&P500 firms. In trade credit supply equations they find that the coefficients of R are again significant and positive but over 50% larger in the model estimate for non-S&P500 firms. Kim and Choi conclude that tighter monetary policy produces an expansion of trade credit – consistent with a trade credit channel – but with trade credit flowing mainly from non-S&P500 firms. We believe two factors explain the difference between their findings and ours. First, our credit market access variables CMA1 and CMA2 (WW-score

[36] Choi and Kim estimate separate supply and demand models of trade credit on quarterly Compustat data for 659 S&P 500 firms and 689 non-S&P 500 firms from the end of 1975 to the end of 1997. Estimates of the models most similar to ours appear in the b columns of their Tables 1 and 2. As we do, they define the supply and demand for trade credit as AR/A_{-1} and AP/A_{-1}, respectively. The explanatory variables in their supply models include $Sales/A_{-1}$ and LnA_{-1} (which also appear in our models), $\Delta Sales^{+}/A_{-1}$ and $\Delta Sales^{-}/A_{-1}$ (positive and negative changes in the sales-to-asset ratio), $[\,LnA_{-1}]^2$, INV/A_{-1}, the retained-earnings-to-assets ratio, and Ln(Short-term $Debt/A_{-1}$). The explanatory variables in their demand models include LnA_{-1} and INV/A_{-1} (which also appear in our models), the cost of goods sold to asset ratio, $\Delta COG^{+}/A_{-1}$ and $\Delta COGS^{-}/A_{-1}$, $[\,LnA_{-1}]^2$, and the retained-earnings-to-assets ratio.

[37] Choi and Kim also use two other measures of policy stance in alternative specifications of their models: change in the federal funds rate and a binary indicator of quarters in which Federal Reserve policy became explicitly disinflationary. The results are qualitatively similar to the results using the 3-month Treasury bill rate.

and Z-score) identify more precisely firms which supply or demand trade credit than membership in a particular stock market index. Second, we allow the impact of tight money to vary by firms' market access through the use of tight money / credit market access interactions terms.[38]

Although our evidence points to a traditional trade credit channel, our estimated models also show evidence of credit flows from constrained firms to unconstrained firms in loose-money periods. In Table 6 the estimated coefficients of the CMA2 quartile indicators in column (3) imply that *constrained* firms in the next-to-poorest, next-to-best and best credit-market-access quartiles provide, respectively, about 5½%, 10¼% and 10½% more accounts receivable in loose money periods than firms with the poorest credit market access; the comparable metrics for the *unconstrained* panel (column 2) are 4%, 10% and 14½%, respectively. The literature offers at least two explanations for extensions of trade credit by financially weak seller firms to financially strong buyer firms. First, weak sellers are often informational opaque, making them risky counterparties. Trade credit permits buyers to inspect goods from opaque sellers and return defective product without loss of cash (Smith, 1987; Long et al., 1993). Second, weak sellers may be coerced by strong buyers into extending trade credit on terms which lower effective product prices (Fabbri and Klapper, 2009; Murjin and Njoroge, 2015). Klapper et al. (2012) find that buyers with market power receive lower effective prices from sellers lacking power who sell

[38] Like the estimated models Choi and Kim report, the estimated models we report in Tables 8 and 9 represent the stance of monetary policy with a market interest rate. In our estimated trade credit supply models (Table 8) the interest rate coefficient is larger in the constrained panel than in the unconstrained panel, similar to Choi and Kim: 0.003 versus 0.001. The estimated coefficients of the tight money/credit market access interaction terms are significantly larger in the unconstrained panel, however, supporting our conclusion that trade credit flows from firms with good credit market access to firms with poorer access.

on credit with long repayment periods. Giannetti et al. (2011) find that buyers with market power receive lower prices through trade credit offered with large early payment discounts. Our empirical findings are consistent with both explanations and add to a growing body of work showing trade credit flows from firms with poor credit market access to firms with good access.

Our empirical findings present two puzzles. First, how can we reconcile the evidence in Table 7 that tight monetary policy has little measurable effect on trade credit demand defined as AP/A, with the evidence in Table 5 that mean AP/A is statistically larger in tight-money periods than in loose-money periods? The most plausible explanation is that tight monetary policy raises mean AP/A through the financial determinants of trade credit demand, namely LnA_{-1}, Tobin's Q_{-1}, $Sales/A_{-1}$, CS/A_{-1}, AR/A_{-1}, INV/A_{-1}, CF/A_{-1}, and LTD/A_{-1}. In the estimated whole sample trade credit demand model (Table 7, column 1), the determinants with the largest coefficients (in absolute value) are LnA_{-1} (inversely related to AP/A), and $Sales/A_{-1}$, AR/A_{-1} and INV/A_{-1} (directly related to AP/A). The coefficient estimates imply elasticities of AP/A with respect to LnA_{-1}, $Sales/A_{-1}$, AR/A_{-1} and INV/A_{-1} of –57.5%, 35.3%, 23% and 20.3%, respectively. Hence a policy tightening might raise AP/A if log-assets fall sufficiently and/or one or more of $Sales/A_{-1}$, AR/A_{-1}, or INV/A_{-1} rises sufficiently. A policy tightening reduces corporate assets, raising AP/A. Nilsen (2002) estimates VAR models in which tighter monetary policy leads to an 8-quarter drop in GDP, reducing Sales and lowering AP/A_{-1}, and an 8-quarter rise in inventories, raising AP/A. The estimated trade credit supply models in Table 6 show that a monetary policy tightening raises AR/A_{-1}, raising AP/A. Thus tighter monetary could plausibly raise AP/A despite statistically insignificant estimated coefficients on the tight money indicators and the tight money / credit market access interaction terms.

Second, how can tight monetary policy expand trade credit supply more than trade credit demand considering that a firm extending trade credit creates an account receivable on its own balance sheet and an account payable of equal size on the balance sheet of the borrowing firm? Since our sample covers essentially the universe of Compustat firms, the most plausible explanation is that the missing accounts payable appear on the balance sheets of private corporations.[39] Evidence presented by Demiroglu et al. (2012) supports this view: their study of private and small public firms shows that tight monetary policy curtails bank lending primarily to private firms which lack established lines of credit and which, therefore, have few financing alternatives aside from trade credit. Evidence by Garcia-Appendini and Montorliol-Garriga (2013) indirectly supports this argument: their study of public supplier-client pairs around the 2008 financial crisis shows that trade credit flowed from suppliers with the greatest liquidity to clients with the poorest credit market access, a result which should generalize to private clients.

V. Conclusion

This paper re-examines the case for a trade credit channel in the US which mitigates impacts from a monetary policy tightening as first described by Meltzer (1960). Our empirical strategy differs from those in previous studies by employing quarterly firm-level data on essentially the universe of publicly-traded corporations over an extended time frame on which we estimate firm-level trade credit supply and demand models. Our strategy also differs from previous studies in using the Whited and Wu (2006) index and the Altman (1968) Z-score index as modified by MacKie-Mason (1990) to represent firms' degree of credit market access. These

[39] Choi and Kim (2005) suggest this possibility to explain their result that tight money expands trade credit supply more than trade credit demand however they cannot prove it since their data do not span the Compustat universe.

metrics permit finer distinctions between degrees of credit market access than possible from firm size or presence of a bond rating, metrics used in prior studies. Distinguishing accurately among degrees of credit market access is critical to uncovering evidence of a trade credit channel since its existence originates from financial market frictions which cause disparities in credit market access among firms.

Our empirical results both confirm the existence of a trade credit channel as conceived by Meltzer (1960) and provide new evidence on its mechanics. In particular, we find that a monetary policy tightening leads to an expansion in the supply of trade credit which is declining in public firms' credit market access, a previously undocumented result. A policy tightening also results in a nearly negligible expansion in the demand for trade credit by public firms, a finding which contrasts with Nilsen (2002) and Choi and Kim (2005) who report evidence of a trade credit channel from smaller firm samples. Together our findings imply that a trade credit channel works by funneling credit from public sector firms commensurate with their credit market access to private sector firms with few financing alternatives, a previously undocumented result.

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Table 1 Variable Definitions

Variable Name	Variable Description			
CMA1 (WW-Score)	Whited and Wu (2006) index. Higher WW-score implies poorer credit market access.			
	WW-score = $-0.091(\text{EBITDA}/\text{A}_{-1})$ -0.062 PDIVD + 0.021 (LTD /A) – 0.044 LnA + 0.120 ISG – 0.035 SG where			
	$\text{EBITDA}/\text{A}_{-1}$	Earnings before interest, taxes and depreciation / lagged assets	PDIVD	Positive cash dividends dummy
	LTD / A	Long-term debt / assets	LnA	Log of inflation-adjusted assets
	ISG	Percentage change in annual inflation-adjusted industry sales, by 3-digit SIC code	SG	Percentage change in annual inflation-adjusted sales
CMA 2 (Z-score)	Altman (1968) Z-score as modified by MacKie-Mason (1990), multiplied by –1. Higher Z-score implies poorer credit market access.			
	Z-score = –3.3 (EBIT/A) – 1.0 (Sales/A) – 1.4 (ARE/A) – 1.2 (NWC/A) where			
	EBIT / A	Earnings after depreciation and before interest and taxes / assets	Sales / A	Sales / assets
	ARE/A	(Accumulated) retained earnings / assets	NWC / A	Net working capital / assets
A	Inflation-adjusted assets			
Age	Firm age, measured as number of years of non-missing asset data in Compustat			
AP/A	Accounts Payable / Assets			
AR/A	Accounts Receivable / Assets			
CF/A	Cash flow-to-assets ratio, measured as (Net Income + Depreciation) / Assets			
CS/A	(Cash + Short Term Securities) / Assets			
FFR	Effective federal funds rate			
INV/A	Inventories / Assets			
Tobin's Q	Market value of assets/book value of assets			
D_CMA2_n	0-1 indicator variable coded 1 if a firm is in CMA2 quartile n. D_CMA2_1=1 (D_CMA2_4=1) indicates a firm in the quartile with the best (poorest) credit market access.			
D_TMP	0-1 indicator variable coded 1 if monetary policy is tight, defined as an end-of-quarter federal funds rate greater or equal to 5%.			

Table 2 Descriptive Statistics

This table presents numbers of observations (NOBS), means, medians and standard deviations (Std Dev) for the variables defined in Table 1. The whole sample comprises the quarterly Compustat universe between 1988 and 2008 excluding financial firms (SIC Code 60-69), utility firms (SIC Code 49), non-classified firms (SIC Code 99), firms with negative or zero assets, and firms with sales less than $5 million. Unconstrained (constrained) firms are firms in the least (most) constrained quartile of the CMA1 distribution in the current quarter and have the best (poorest) credit market access. CMA1 is the Whited and Wu (2006) index defined in Table 1.

Variable	Whole Sample				Unconstrained Firms				Constrained Firms			
	NOBS	Mean	Median	Std Dev	NOBS	Mean	Median	Std Dev	NOBS	Mean	Median	Std Dev
Assets†	378244	2690	310	12634	72527	5863	923	19383	72527	1127	145	7864
Sales†	378244	624	80	2833	72527	1322	216	4181	72527	280	38	2460
CMA1 (WW-score)	303633	-0.55	-0.56	0.34	70873	-0.83	-0.80	0.28	70832	-0.26	-0.28	0.31
CMA2 (Z-score)	343193	-0.60	-0.81	1.28	65046	-0.71	-0.80	0.98	67485	-0.31	-0.68	1.64
AR/A	366681	0.18	0.16	0.13	70201	0.16	0.14	0.12	71113	0.20	0.18	0.14
AP/A	371196	0.10	0.08	0.09	71165	0.10	0.08	0.08	71833	0.11	0.08	0.09
Age ‡	378244	11.39	9.00	8.51	72527	13.33	11.00	9.36	72527	11.03	9.00	7.59
ARE/A	355769	-0.05	0.11	0.80	67905	0.07	0.15	0.58	69657	-0.24	0.05	1.04
CF/A	378198	0.01	0.02	0.05	72522	0.02	0.02	0.04	72523	0.01	0.02	0.06
CS/A	374232	0.13	0.06	0.17	71712	0.11	0.04	0.16	72067	0.15	0.07	0.18
EBIT/A	376356	0.02	0.02	0.04	72236	0.02	0.02	0.03	72191	0.01	0.01	0.04
EBITDA/A	342522	0.03	0.03	0.04	71607	0.04	0.03	0.03	71432	0.03	0.03	0.04
INV/A	310214	0.18	0.15	0.15	61645	0.16	0.12	0.14	56295	0.19	0.15	0.17
ISG	368529	1.54	1.48	2.60	72527	0.29	0.54	2.62	72527	2.94	2.45	2.95
LnA	378244	14.38	14.18	1.78	72527	15.38	15.27	1.82	72527	13.66	13.42	1.52
LTD/A	373945	0.21	0.16	0.22	72301	0.23	0.20	0.21	72283	0.19	0.11	0.22
NWC	365601	0.23	0.21	0.24	69591	0.19	0.16	0.21	70445	0.24	0.24	0.26
PDIVD	367962	0.36	0.00	0.48	71909	0.51	1.00	0.50	72048	0.22	0.00	0.41
Sales/A	378244	0.33	0.28	0.22	72527	0.30	0.26	0.21	72527	0.34	0.29	0.23
SG	342168	1.97	1.29	7.34	72527	4.56	2.55	8.25	72527	-0.11	0.05	7.58
Tobin's Q	332791	1.86	1.42	1.47	64591	1.92	1.46	1.52	67173	1.69	1.34	1.15

† millions of 2008 dollars ‡ years

Table 3 Descriptive Statistics for Accounts Receivable and Accounts Payable to Assets, by CMA1 Deciles

This table presents numbers of observations (NOBS), means, medians and standard deviations (Std Dev) for accounts receivable and accounts payable to assets by CMA1 decile. CMA1 is the index of Whited and Wu (2006) defined in Table 1. Firms in Decile 1 (10) are the least (most) financially constrained and have the best (poorest) credit market access. The sample comprises the quarterly Compustat universe of industrial firms between 1988 and 2008 excluding financial firms (SIC Code 60-69), utility firms (SIC Code 49), non-classified firms (SIC Code 99), firms with negative or zero assets, and firms with sales less than $5 million.

CMA1 Decile	Decile Mean CMA1	Accounts Receivable (AR/A)				Accounts Payable (AP/A)			
		NOBS	Mean	Median	Std Dev	NOBS	Mean	Median	Std Dev
Best access 1	-1.15	29415	0.16	0.13	0.12	29907	0.10	0.07	0.08
2	-0.84	29524	0.16	0.14	0.12	29877	0.10	0.08	0.08
3	-0.73	29565	0.16	0.14	0.12	29903	0.10	0.08	0.08
4	-0.66	29539	0.17	0.15	0.12	29963	0.10	0.08	0.08
5	-0.59	29624	0.17	0.15	0.12	29988	0.10	0.08	0.08
6	-0.53	29662	0.18	0.16	0.13	30008	0.10	0.08	0.09
7	-0.45	29710	0.18	0.17	0.13	30061	0.10	0.08	0.09
8	-0.37	29726	0.19	0.18	0.13	30069	0.11	0.08	0.09
9	-0.25	29737	0.20	0.18	0.14	30079	0.11	0.08	0.09
Poorest access 10	0.08	29740	0.20	0.17	0.14	30058	0.11	0.08	0.09
All		296242	0.18	0.16	0.13	299913	0.10	0.08	0.09

Table 4 Mean Accounts Receivable to Assets, Accounts Payable to Assets, CMA1 and CMA2, by Year

This table presents sample mean values of accounts receivable to assets (AR/A), accounts payable to assets (AP/A), and measures of credit market access CMA1 (the index of Whited and Wu, 2006), and CMA2 (Z-score of Altman, 1968, as modified by MacKie-Mason, 1990, multiplied by -1). Higher values of CMA1 and CMA2 imply poorer credit market access. The whole sample (WS) comprises the quarterly Compustat universe of industrial firms between 1988 and 2008 excluding financial firms (SIC Code 60-69), utility firms (SIC Code 49), non-classified firms (SIC Code 99), firms with negative or zero assets, and firms with sales less than $5 million. The unconstrained (U) and constrained (C) subsamples comprise firms in the least (most) constrained quartiles of the CMA1 distribution in the current quarter and have the best (poorest) credit market access.

	AR/A			AP/A			CMA1			CMA2		
Year	WS	U	C	WS	U	C	WS	U	C	WS	U	C
1988	0.21	0.17	0.23	0.11	0.10	0.12	-0.51	-0.85	-0.17	-1.05	-1.00	-0.96
1989	0.21	0.18	0.23	0.11	0.10	0.12	-0.60	-0.92	-0.28	-1.01	-0.96	-0.88
1990	0.20	0.19	0.22	0.11	0.10	0.12	-0.63	-0.96	-0.29	-0.96	-1.00	-0.78
1991	0.20	0.18	0.22	0.11	0.10	0.13	-0.70	-1.03	-0.36	-0.93	-0.95	-0.79
1992	0.20	0.16	0.23	0.11	0.10	0.12	-0.59	-0.91	-0.26	-0.95	-0.89	-0.87
1993	0.20	0.16	0.22	0.11	0.10	0.12	-0.57	-0.90	-0.22	-0.92	-0.95	-0.80
1994	0.20	0.16	0.23	0.11	0.11	0.12	-0.47	-0.82	-0.11	-0.91	-0.90	-0.80
1995	0.20	0.16	0.24	0.11	0.10	0.12	-0.47	-0.83	-0.09	-0.89	-0.91	-0.76
1996	0.20	0.17	0.22	0.11	0.11	0.11	-0.50	-0.88	-0.09	-0.87	-0.94	-0.68
1997	0.20	0.17	0.23	0.11	0.10	0.11	-0.47	-0.88	-0.04	-0.82	-0.91	-0.60
1998	0.19	0.16	0.23	0.10	0.10	0.11	-0.50	-0.96	-0.04	-0.73	-0.77	-0.54
1999	0.19	0.16	0.20	0.10	0.09	0.11	-0.51	-0.94	-0.07	-0.64	-0.73	-0.42
2000	0.18	0.16	0.20	0.10	0.10	0.10	-0.48	-0.95	0.06	-0.55	-0.70	-0.34
2001	0.16	0.16	0.17	0.10	0.10	0.10	-0.72	-1.20	-0.26	-0.33	-0.54	0.10
2002	0.16	0.16	0.16	0.10	0.09	0.10	-0.67	-1.08	-0.27	-0.23	-0.49	0.23
2003	0.16	0.16	0.16	0.10	0.10	0.10	-0.55	-0.88	-0.19	-0.26	-0.61	0.20
2004	0.16	0.14	0.17	0.10	0.09	0.10	-0.46	-0.80	-0.10	-0.37	-0.63	-0.03
2005	0.16	0.14	0.17	0.10	0.09	0.10	-0.53	-0.87	-0.15	-0.36	-0.63	0.05
2006	0.16	0.14	0.18	0.10	0.09	0.10	-0.52	-0.87	-0.15	-0.38	-0.65	0.13
2007	0.15	0.14	0.17	0.09	0.10	0.10	-0.55	-0.91	-0.19	-0.38	-0.71	0.19
2008	0.15	0.14	0.16	0.09	0.10	0.09	-0.60	-1.02	-0.14	-0.32	-0.65	0.16

Table 5 Descriptive Statistics for CMA1, Accounts Receivable to Assets and Accounts Payable to Assets, by Monetary Policy Stance

This table presents statistics on CMA1 (the index of Whited and Wu, 2006), accounts receivable to assets (AR/A) and accounts payable to assets (AP/A) by monetary policy stance. Stance is measured by the end-of-quarter federal funds rate: stance is tight (loose) in quarters when the fed funds rate is at least (is less than) 5%. The whole sample comprises the quarterly Compustat universe of industrial firms between 1988 and 2008 excluding financial firms (SIC Code 60-69), utility firms (SIC Code 49), non-classified firms (SIC Code 99), firms with negative or zero assets, and firms with sales less than $5 million. Panel A shows descriptive statistics for CMA1 by policy stance. Higher CMA1 values imply poorer credit market access. Panel B shows mean AR/A and AP/A for the whole sample (WS) and for the unconstrained (U) and constrained (C) subsamples. Unconstrained (constrained) firms fall into the least (most) constrained quartiles of the CMA1 distribution in the current quarter and have the best (poorest) credit market access.

Panel A: Descriptive Statistics for CMA1 by Monetary Policy Stance for the Whole Sample

Statistic:	NOBS	Mean	Median	Std Dev	Min	Max
Monetary Policy Stance:						
Loose	153525	-0.57	-0.57	0.33	-3.26	2.91
Tight	150108	-0.53*	-0.54	0.34	-2.69	3.09
All	303633	-0.55	-0.56	0.34	-3.26	3.09

Panel B: Mean AR/A and AP/A for the Whole Sample and Subsamples, by Monetary Policy Stance

	AR/A			AP/A		
Sample:	WS	U	C	WS	U	C
Monetary Policy Stance:						
Loose	0.171	0.151	0.185†	0.100	0.095	0.106†
Tight	0.190‡	0.163‡	0.215†‡	0.106‡	0.099‡	0.111†‡
All	0.179	0.157	0.200†	0.103	0.097	0.109†

* The statistic for tight policy is significantly greater than the statistic for loose policy.
† The statistic for constrained firms is significantly greater than the statistic for unconstrained firms in the same policy stance.
‡ The statistic for the tight policy stance is significantly greater than the statistic for the loose policy stance for firms in the same subsample.

Table 6 Estimated Supply of Trade Credit Models

This table presents estimates of equation 1. The dependent variable is accounts receivable to assets, AR/A. D_CMA2_n is a 0-1 variable coded 1 if a firm is in CMA2 quartile n; D_CMA2_4, the quartile with the poorest credit market access, is the excluded category. CMA2 is the Altman (1968) measure of financial distress as modified by MacKie-Mason (1990), multiplied by –1. D_TMP is a 0-1 variable coded 1 in a tight-money period (the end-of-quarter fed funds rate is at least 5%). LnA is the natural log of assets, LnAge is the natural log of firm age, Tobin's Q is the market to book ratio for assets, Sales/A is sales to assets, CS/A is cash and short-term securities to assets, and AP/A is accounts payable to assets. WS denotes the whole sample and U (C) denotes the unconstrained (constrained) subsample with the best (poorest) credit market access. The whole sample comprises the quarterly Compustat universe of industrial firms between 1988 and 2008. Firms in U (C) fall into the first (last) quartile of the CMA1 distribution. CMA1 is the Whited and Wu (2006) index. Coefficient estimates (robust standard errors) appear in the first (second) rows for each independent variable. ***, **, and * denote statistical significance at the 1%-, 5%- and 10%-levels, respectively. Each estimated model includes unreported quarterly dummies and firm fixed-effects constants. All variables are winsorized at 1% from their lower and upper bounds.

Sample:		WS (1)	U (2)	C (3)
Independent Variables (predicted coefficients):				
$D\_CMA2_{1,t-1}$	$(\beta_1^s > \beta_2^s > \beta_3^s > 0)$†	0.021***	0.022***	0.020***
		(0.002)	(0.003)	(0.003)
$D\_CMA2_{2,t-1}$	$(\beta_1^s > \beta_2^s > \beta_3^s > 0)$†	0.016***	0.015***	0.019***
		(0.001)	(0.002)	(0.003)
$D\_CMA2_{3,t-1}$	$(\beta_1^s > \beta_2^s > \beta_3^s > 0)$†	0.008***	0.006***	0.010***
		(0.001)	(0.001)	(0.002)
D_TMP_{t-1}	$(\varphi^s > 0)$‡	0.006***	0.004***	0.004
		(0.001)	(0.001)	(0.002)
$D\_TMP_{t-1} * D\_CMA2_{1,t-1}$	$(\gamma_1^s > \gamma_2^s > \gamma_3^s > 0)$§	0.005***	0.009***	0.001
		(0.001)	(0.002)	(0.002)
$D\_TMP_{t-1} * D\_CMA2_{2,t-1}$	$(\gamma_1^s > \gamma_2^s > \gamma_3^s > 0)$§	0.005***	0.008***	-0.001
		(0.001)	(0.002)	(0.002)
$D\_TMP_{t-1} * D\_CMA2_{3,t-1}$	$(\gamma_1^s > \gamma_2^s > \gamma_3^s > 0)$§	0.003***	0.006***	-0.002
		(0.001)	(0.001)	(0.002)
LnA_{t-1}	$(\delta_1^s < 0)$	-0.016***	-0.014***	-0.016***
		(0.001)	(0.001)	(0.002)
$LnAge_{t-1}$	$(\delta_2^s < 0)$	-0.006***	-0.005***	-0.008***
		(0.001)	(0.002)	(0.002)
$Tobin's\ Q_{t-1}$	$(\delta_3^s > 0)$	0.001***	0.001***	0.003***
		(0.000)	(0.000)	(0.001)
$Sales/A_{t-1}$	$(\delta_4^s > 0)$	0.122***	0.088***	0.140***
		(0.005)	(0.008)	(0.009)
CS/A_{t-1}	$(\delta_5^s < 0)$	-0.129***	-0.107***	-0.144***
		(0.004)	(0.006)	(0.006)
AP/A_{t-1}	$(\delta_6^s > 0)$	0.207***	0.234***	0.179***
		(0.010)	(0.016)	(0.015)
$D\_TMP_{t-1} * Sales/A_{t-1}$	$(\delta_7^s = ?)$	-0.013***	-0.017***	-0.000
		(0.003)	(0.004)	(0.005)
$D\_TMP_{t-1} * CS/A_{t-1}$	$(\delta_8^s = ?)$	-0.001	0.004	0.000
		(0.002)	(0.004)	(0.004)
Number of Observations		288,430	58,408	62,717
Number of Firms		10,966	7,231	7,789
Adj. R-squared		0.23	0.22	0.22

† β_n^s greater for unconstrained firms than constrained firms. ‡ φ^s greater for unconstrained firms than constrained firms.
§ γ_n^s greater for unconstrained firms than constrained firms.

Table 7 Estimated Demand for Trade Credit Models

This table presents estimates of equation 2. The dependent variable is accounts payable to assets, AP/A. D_CMA2_n is a 0-1 variable coded 1 if a firm is in CMA2 quartile n; D_CMA2_4, the quartile with the poorest credit market access, is the excluded category. CMA2 is the Altman (1968) measure of financial distress as modified by MacKie-Mason (1990), multiplied by –1. D_TMP is a 0-1 variable coded 1 in a tight-money period (the end-of-quarter fed funds rate is at least 5%). LnA is the natural log of assets, LnAge is the natural log of firm age, Tobin's Q is the market to book ratio for assets, Sales/A is sales to assets, CS/A is cash and short-term securities to assets, AR/A is accounts receivable to assets, INV/A is inventory to assets, CF/A is cash flow to assets, and LTD/A is long-term debt to assets. WS denotes the whole sample and U (C) denotes the unconstrained (constrained) subsample with the best (poorest) credit market access. The whole sample comprises the quarterly Compustat universe of industrial firms between 1988 and 2008. Firms in U (C) fall into the first (last) quartile of the CMA1 distribution. CMA1 is the Whited and Wu (2006) index. Coefficient estimates (robust standard errors) appear in the first (second) rows for each independent variable. ***, **, and * denote statistical significance at the 1%-, 5%- and 10%-levels, respectively. Each estimated model includes unreported quarterly dummies and firm fixed-effects constants. All variables are winsorized at 1% from their lower and upper bounds.

Sample:		WS (1)	U (2)	C (3)
Independent Variables (predicted coefficients):				
$D\_CMA2_{1,t-1}$	$(\beta_1^d < \beta_2^d < \beta_3^d < 0)$ †	-0.028***	-0.019***	-0.035***
		(0.002)	(0.002)	(0.003)
$D\_CMA2_{2,t-1}$	$(\beta_1^d < \beta_2^d < \beta_3^d < 0)$ †	-0.018***	-0.011***	-0.024***
		(0.001)	(0.002)	(0.002)
$D\_CMA2_{3,t-1}$	$(\beta_1^d < \beta_2^d < \beta_3^d < 0)$ †	-0.010***	-0.006***	-0.014***
		(0.001)	(0.001)	(0.002)
D_TMP_{t-1}	$(\varphi^d > 0)$ ‡	0.001	0.001	-0.002
		(0.001)	(0.001)	(0.002)
$D\_TMP_{t-1} * D\_CMA2_{1,t-1}$	$(\gamma_1^d < \gamma_2^d < \gamma_3^d < 0)$ §	-0.003**	-0.004**	-0.002
		(0.001)	(0.002)	(0.002)
$D\_TMP_{t-1} * D\_CMA2_{2,t-1}$	$(\gamma_1^d < \gamma_2^d < \gamma_3^d < 0)$ §	-0.001	-0.002	-0.000
		(0.001)	(0.002)	(0.002)
$D\_TMP_{t-1} * D\_CMA2_{3,t-1}$	$(\gamma_1^d < \gamma_2^d < \gamma_3^d < 0)$ §	0.000	0.000	-0.000
		(0.001)	(0.001)	(0.002)
LnA_{t-1}	$(\delta_1^d < 0)$	-0.004***	-0.004***	-0.006***
		(0.001)	(0.001)	(0.002)
$LnAge_{t-1}$	$(\delta_2^d < 0)$	0.007***	0.006***	0.009***
		(0.001)	(0.001)	(0.002)
$Tobin's\ Q_{t-1}$	$(\delta_3^d > 0)$	0.001***	0.001**	0.002***
		(0.000)	(0.000)	(0.000)
$Sales/A_{t-1}$	$(\delta_4^d > 0)$	0.107***	0.102***	0.104***
		(0.005)	(0.007)	(0.008)
CS/A_{t-1}	$(\delta_5^d > 0)$	0.007**	0.014***	-0.011*
		(0.003)	(0.005)	(0.006)
AR/A_{t-1}	$(\delta_6^d > 0)$	0.128***	0.121***	0.125***
		(0.007)	(0.011)	(0.011)
INV/A_{t-1}	$(\delta_7^d > 0)$	0.113***	0.109***	0.121***
		(0.008)	(0.012)	(0.012)
CF/A_{t-1}	$(\delta_8^d < 0)$	-0.085***	-0.082***	-0.072***
		(0.006)	(0.011)	(0.010)
LTD/A_{t-1}	$(\delta_9^d < 0)$	-0.005**	-0.007*	0.001
		(0.002)	(0.004)	(0.004)
$D\_TMP_{t-1} * Sales/A_{t-1}$	$(\delta_{10}^d = ?)$	-0.002	0.002	0.001
		(0.002)	(0.004)	(0.004)
$D\_TMP_{t-1} * CS/A_{t-1}$	$(\delta_{11}^d = ?)$	0.007***	0.006*	0.008*
		(0.002)	(0.003)	(0.004)
$D\_TMP_{t-1} * CF/A_{t-1}$	$(\delta_{12}^d = ?)$	-0.016**	-0.018	-0.008
		(0.008)	(0.017)	(0.015)
Number of Observations		238,850	50,416	49,202
Number of Firms		9,074	5,996	6,369
Adj. R-squared		0.17	0.17	0.17

† $|\beta_n^d|$ greater for constrained firms than unconstrained firms. ‡ φ^d greater for constrained firms than unconstrained firms.
§ $|\gamma_n^d|$ greater for constrained firms than unconstrained firms.

Table 8 Robustness Check: Estimated Supply of Trade Credit Models with the Federal Funds Rate

This table presents estimates of equation 1. The dependent variable is accounts receivable to assets, AR/A. D_CMA2_n is a 0-1 variable coded 1 if a firm is in CMA2 quartile n; D_CMA2_4, the quartile with the poorest credit market access, is the excluded category. CMA2 is the Altman (1968) measure of financial distress as modified by MacKie-Mason (1990), multiplied by –1. FFR is the federal funds rate. LnA is the natural log of assets, LnAge is the natural log of firm age, Tobin's Q is the market to book ratio for assets, Sales/A is sales to assets, CS/A is cash and short-term securities to assets, and AP/A is accounts payable to assets. WS denotes the whole sample and U (C) denotes the unconstrained (constrained) subsample with the best (poorest) credit market access. The whole sample comprises the quarterly Compustat universe of industrial firms between 1988 and 2008. Firms in U (C) fall into the first (last) quartile of the CMA1 distribution. CMA1 is the Whited and Wu (2006) index. Coefficient estimates (robust standard errors) appear in the first (second) rows for each independent variable. ***, **, and * denote statistical significance at the 1%-, 5%- and 10%-levels, respectively. Each estimated model includes unreported quarterly dummies and firm fixed-effects constants. All variables are winsorized at 1% from their lower and upper bounds.

Sample:		WS (1)	U (2)	C (3)
Independent Variables (predicted coefficients):				
$D\_CMA2_{1,t-1}$	$(\beta_1^s > \beta_2^s > \beta_3^s > 0)$†	0.016***	0.003***	0.019***
		(0.002)	(0.001)	(0.004)
$D\_CMA2_{2,t-1}$	$(\beta_1^s > \beta_2^s > \beta_3^s > 0)$†	0.010***	0.003***	0.019***
		(0.002)	(0.001)	(0.004)
$D\_CMA2_{3,t-1}$	$(\beta_1^s > \beta_2^s > \beta_3^s > 0)$†	0.006***	0.002***	0.012***
		(0.002)	(0.000)	(0.003)
FFR_{t-1}	$(\varphi^s > 0)$‡	0.002***	0.001***	0.003***
		(0.000)	(0.000)	(0.001)
$FFR_{t-1} * D\_CMA2_{1,t-1}$	$(\gamma_1^s > \gamma_2^s > \gamma_3^s > 0)$§	0.002***	0.010***	0.000
		(0.000)	(0.004)	(0.001)
$FFR_{t-1} * D\_CMA2_{2,t-1}$	$(\gamma_1^s > \gamma_2^s > \gamma_3^s > 0)$§	0.002***	0.005*	-0.000
		(0.000)	(0.003)	(0.001)
$FFR_{t-1} * D\_CMA2_{3,t-1}$	$(\gamma_1^s > \gamma_2^s > \gamma_3^s > 0)$§	0.001**	-0.001	-0.001
		(0.000)	(0.002)	(0.001)
LnA_{t-1}	$(\delta_1^s < 0)$	-0.016***	-0.014***	-0.016***
		(0.001)	(0.001)	(0.002)
$LnAge_{t-1}$	$(\delta_2^s < 0)$	-0.004***	-0.003*	-0.006***
		(0.001)	(0.002)	(0.002)
$Tobin's\ Q_{t-1}$	$(\delta_3^s > 0)$	0.001***	0.001***	0.003***
		(0.000)	(0.000)	(0.001)
$Sales/A_{t-1}$	$(\delta_4^s > 0)$	0.141***	0.107***	0.152***
		(0.007)	(0.009)	(0.011)
CS/A_{t-1}	$(\delta_5^s < 0)$	-0.124***	-0.108***	-0.137***
		(0.005)	(0.008)	(0.008)
AP/A_{t-1}	$(\delta_6^s > 0)$	0.204***	0.233***	0.179***
		(0.010)	(0.016)	(0.015)
$FFR_{t-1} * Sales/A_{t-1}$	$(\delta_7^s = ?)$	-0.005***	-0.006***	-0.002
		(0.001)	(0.001)	(0.002)
$FFR_{t-1} * CS/A_{t-1}$	$(\delta_8^s = ?)$	-0.001	0.001	-0.001
		(0.001)	(0.001)	(0.001)
Number of Observations		264,827	58,411	62,733
Number of Firms		10,547	7,231	7,789
Adj. R-squared		0.23	0.22	0.22

† β_n^s greater for unconstrained firms than constrained firms. ‡ φ^s greater for unconstrained firms than constrained firms.
§ γ_n^s greater for unconstrained firms than constrained firms.

Table 9 Robustness Check: Estimated Demand for Trade Credit Models with the Federal Funds Rate

This table presents estimates of equation 2. The dependent variable is accounts payable to assets, AP/A. D_CMA2_n is a 0-1 variable coded 1 if a firm is in CMA2 quartile n; D_CMA2_4, the quartile with the poorest credit market access, is the excluded category. CMA2 is the Altman (1968) measure of financial distress as modified by MacKie-Mason (1990), multiplied by –1. FFR is the federal funds rate. LnA is the natural log of assets, LnAge is the natural log of firm age, Tobin's Q is the market to book ratio for assets, Sales/A is sales to assets, CS/A is cash and short-term securities to assets, AR/A is accounts receivable to assets, INV/A is inventory to assets, CF/A is cash flow to assets, and LTD/A is long-term debt to assets. WS denotes the whole sample and U (C) denotes the unconstrained (constrained) subsample with the best (poorest) credit market access. The whole sample comprises the quarterly Compustat universe of industrial firms between 1988 and 2008. Firms in U (C) fall into the first (last) quartile of the CMA1 distribution. CMA1 is the Whited and Wu (2006) index. Coefficient estimates (robust standard errors) appear in the first (second) rows for each independent variable. ***, **, and * denote statistical significance at the 1%-, 5%- and 10%-levels, respectively. Each estimated model includes unreported quarterly dummies and firm fixed-effects constants. All variables are winsorized at 1% from their lower and upper bounds.

Sample:		WS (1)	U (2)	C (3)
Independent Variables (predicted coefficients):				
$D\_CMA2_{1,t-1}$	$(\beta_1^d < \beta_2^d < \beta_3^d < 0)$ †	-0.025***	-0.014***	-0.033***
		(0.002)	(0.003)	(0.004)
$D\_CMA2_{2,t-1}$	$(\beta_1^d < \beta_2^d < \beta_3^d < 0)$ †	-0.016***	-0.007***	-0.022***
		(0.002)	(0.003)	(0.004)
$D\_CMA2_{3,t-1}$	$(\beta_1^d < \beta_2^d < \beta_3^d < 0)$ †	-0.010***	-0.005**	-0.013***
		(0.002)	(0.002)	(0.003)
FFR_{t-1}	$(\varphi^d > 0)$ ‡	0.000	0.001	-0.001
		(0.000)	(0.000)	(0.001)
$FFR_{t-1} * D\_CMA2_{1,t-1}$	$(\gamma_1^d < \gamma_2^d < \gamma_3^d < 0)$ §	-0.001***	-0.002***	-0.001
		(0.000)	(0.001)	(0.001)
$FFR_{t-1} * D\_CMA2_{2,t-1}$	$(\gamma_1^d < \gamma_2^d < \gamma_3^d < 0)$ §	-0.001*	-0.001**	-0.000
		(0.000)	(0.000)	(0.001)
$FFR_{t-1} * D\_CMA2_{3,t-1}$	$(\gamma_1^d < \gamma_2^d < \gamma_3^d < 0)$ §	-0.000	-0.000	-0.000
		(0.000)	(0.000)	(0.001)
LnA_{t-1}	$(\delta_1^d < 0)$	-0.004***	-0.004***	-0.006***
		(0.001)	(0.001)	(0.002)
$LnAge_{t-1}$	$(\delta_2^d < 0)$	0.007***	0.007***	0.009***
		(0.001)	(0.001)	(0.002)
Tobin's Q_{t-1}	$(\delta_3^d > 0)$	0.001***	0.001**	0.001***
		(0.000)	(0.000)	(0.000)
$Sales/A_{t-1}$	$(\delta_4^d > 0)$	0.109***	0.101***	0.103***
		(0.006)	(0.010)	(0.009)
CS/A_{t-1}	$(\delta_5^d > 0)$	-0.002	0.010	-0.022***
		(0.004)	(0.007)	(0.008)
AR/A_{t-1}	$(\delta_6^d > 0)$	0.126***	0.120***	0.125***
		(0.007)	(0.011)	(0.011)
INV/A_{t-1}	$(\delta_7^d > 0)$	0.113***	0.110***	0.121***
		(0.007)	(0.012)	(0.012)
CF/A_{t-1}	$(\delta_8^d < 0)$	-0.086***	-0.078***	-0.087***
		(0.011)	(0.018)	(0.018)
LTD/A_{t-1}	$(\delta_9^d < 0)$	-0.005**	-0.007*	0.001
		(0.002)	(0.004)	(0.004)
$D\_TMP_{t-1} * Sales/A_{t-1}$	$(\delta_{10}^d = ?)$	0.000	0.001	0.001
		(0.001)	(0.001)	(0.001)
$D\_TMP_{t-1} * CS/A_{t-1}$	$(\delta_{11}^d = ?)$	0.002***	0.001	0.003***
		(0.001)	(0.001)	(0.001)
$D\_TMP_{t-1} * CF/A_{t-1}$	$(\delta_{12}^d = ?)$	-0.001	-0.002	0.003
		(0.002)	(0.004)	(0.004)
Number of Observations		218,046	50,419	49,214
Number of Firms		8,698	5,996	6,369
Adj. R-squared		0.17	0.17	0.17

† $|\beta_n^d|$ greater for constrained firms than unconstrained firms. ‡ φ^d greater for constrained firms than unconstrained firms.
§ $|\gamma_n^d|$ greater for constrained firms than unconstrained firms.

Table 10 Robustness Check: Estimated Supply of Trade Credit Models with CMA2 as a Continuous Variable

This table presents estimates of equation 1. The dependent variable is accounts receivable to assets, AR/A. CMA2 is the Z-score of Altman (1968) as modified by MacKie-Mason (1990), multiplied by -1; higher values imply poorer credit market access. D_TMP is a 0-1 variable coded 1 in a tight-money period (the end-of-quarter fed funds rate is at least 5%). LnA is the natural log of assets, LnAge is the natural log of firm age, Tobin's Q is the market to book ratio for assets, Sales/A is sales to assets, CS/A is cash and short-term securities to assets, and AP/A is accounts payable to assets. WS denotes the whole sample and U (C) denotes the unconstrained (constrained) subsample with the best (poorest) credit market access. The whole sample comprises the quarterly Compustat universe of industrial firms between 1988 and 2008. Firms in U (C) fall into the first (last) quartile of the CMA1 distribution. CMA1 is the Whited and Wu (2006) index. Coefficient estimates (robust standard errors) appear in the first (second) rows for each independent variable. ***, **, and * denote statistical significance at the 1%-, 5%- and 10%-levels, respectively. Each estimated model includes unreported quarterly dummies and firm fixed-effects constants. All variables are winsorized at 1% from their lower and upper bounds.

Sample:		WS (1)	U (2)	C (3)
Independent Variables (predicted coefficients):				
$\mathrm{CMA2}_{t-1}$	$(\beta^s < 0)$†	-0.004***	-0.006***	-0.005***
		(0.001)	(0.001)	(0.001)
$\mathrm{D\_TMP}_{t-1}$	$(\varphi^s > 0)$‡	0.007***	0.005***	0.002
		(0.001)	(0.001)	(0.002)
$\mathrm{D\_TMP}_{t-1} * \mathrm{CMA2}_{t-1}$	$(\gamma^s < 0)$§	-0.003***	-0.007***	-0.001*
		(0.000)	(0.001)	(0.001)
LnA_{t-1}	$(\delta_1^s < 0)$	-0.016***	-0.015***	-0.017***
		(0.001)	(0.001)	(0.002)
LnAge_{t-1}	$(\delta_2^s < 0)$	-0.003***	-0.002	-0.005**
		(0.001)	(0.002)	(0.002)
$\mathrm{Tobin's\ Q}_{t-1}$	$(\delta_3^s > 0)$	0.001***	0.001***	0.003***
		(0.000)	(0.000)	(0.001)
$\mathrm{Sales/A}_{t-1}$	$(\delta_4^s > 0)$	0.130***	0.094***	0.144***
		(0.005)	(0.007)	(0.009)
$\mathrm{CS/A}_{t-1}$	$(\delta_5^s < 0)$	-0.124***	-0.102***	-0.142***
		(0.004)	(0.006)	(0.006)
$\mathrm{AP/A}_{t-1}$	$(\delta_6^s > 0)$	0.206***	0.237***	0.180***
		(0.010)	(0.016)	(0.015)
$\mathrm{D\_TMP}_{t-1} * \mathrm{Sales/A}_{t-1}$	$(\delta_7^s = ?)$	-0.013***	-0.020***	0.001
		(0.002)	(0.004)	(0.004)
$\mathrm{D\_TMP}_{t-1} * \mathrm{CS/A}_{t-1}$	$(\delta_8^s = ?)$	-0.001	0.002	0.002
		(0.002)	(0.004)	(0.004)
Number of Observations		288,430	58,408	62,717
Number of Firms		10,966	7,231	7,789
Adj. R-squared		0.23	0.22	0.22

† $|\beta_n^s|$ greater for unconstrained firms than constrained firms. ‡ φ^s greater for unconstrained firms than constrained firms.
§ $|\gamma_n^s|$ greater for unconstrained firms than constrained firms.

Table 11 Robustness Check: Estimated Demand for Trade Credit Models with CMA2 as a Continuous Variable

This table presents estimates of equation 2. The dependent variable is accounts payable to assets, AP/A. CMA2 is the Z-score of Altman (1968) as modified by MacKie-Mason (1990), multiplied by -1; higher values imply poorer credit market access. D_TMP is a 0-1 variable coded 1 in a tight-money period (the end-of-quarter fed funds rate is at least 5%). LnA is the natural log of assets, LnAge is the natural log of firm age, Tobin's Q is the market to book ratio for assets, Sales/A is sales to assets, CS/A is cash and short-term securities to assets, AR/A is accounts receivable to assets, INV/A is inventory to assets, CF/A is cash flow to assets, and LTD/A is long-term debt to assets. WS denotes the whole sample and U (C) denotes the unconstrained (constrained) subsample with the best (poorest) credit market access. The whole sample comprises the quarterly Compustat universe of industrial firms between 1988 and 2008. Firms in U (C) fall into the first (last) quartile of the CMA1 distribution. CMA1 is the Whited and Wu (2006) index. Coefficient estimates (robust standard errors) appear in the first (second) rows for each independent variable. ***, **, and * denote statistical significance at the 1%-, 5%- and 10%-levels, respectively. Each estimated model includes unreported quarterly dummies and firm fixed-effects constants. All variables are winsorized at 1% from their lower and upper bounds.

Sample:		WS (1)	U (2)	C (3)
Independent Variables (predicted coefficients):				
$CMA2_{t-1}$	$(\beta^d > 0)$ †	0.010***	0.008***	0.011***
		(0.001)	(0.001)	(0.001)
D_TMP_{t-1}	$(\varphi^d > 0)$ ‡	0.001	0.002	-0.001
		(0.001)	(0.001)	(0.002)
$D\_TMP_{t-1} * CMA2_{t-1}$	$(\gamma^d > 0)$ §	0.003***	0.005***	0.002***
		(0.000)	(0.001)	(0.001)
LnA_{t-1}	$(\delta_1^d < 0)$	-0.004***	-0.002	-0.003**
		(0.001)	(0.001)	(0.002)
$LnAge_{t-1}$	$(\delta_2^d < 0)$	0.007***	0.003**	0.002
		(0.001)	(0.001)	(0.002)
$Tobin's\ Q_{t-1}$	$(\delta_3^d > 0)$	0.001***	0.001*	0.001**
		(0.000)	(0.000)	(0.000)
$Sales/A_{t-1}$	$(\delta_4^d > 0)$	0.110***	0.100***	0.097***
		(0.005)	(0.007)	(0.007)
CS/A_{t-1}	$(\delta_5^d > 0)$	0.005*	0.013**	-0.014**
		(0.003)	(0.005)	(0.006)
AR/A_{t-1}	$(\delta_6^d > 0)$	0.125***	0.121***	0.123***
		(0.007)	(0.011)	(0.011)
INV/A_{t-1}	$(\delta_7^d > 0)$	0.112***	0.110***	0.120***
		(0.007)	(0.012)	(0.012)
CF/A_{t-1}	$(\delta_8^d < 0)$	-0.085***	-0.074***	-0.066***
		(0.006)	(0.011)	(0.011)
LTD/A_{t-1}	$(\delta_9^d < 0)$	-0.005**	-0.008**	0.000
		(0.002)	(0.004)	(0.004)
$D\_TMP_{t-1} * Sales/A_{t-1}$	$(\delta_{10}^d = ?)$	-0.001	0.002	0.001
		(0.002)	(0.004)	(0.004)
$D\_TMP_{t-1} * CS/A_{t-1}$	$(\delta_{11}^d = ?)$	0.006***	0.006*	0.007*
		(0.002)	(0.003)	(0.004)
$D\_TMP_{t-1} * CF/A_{t-1}$	$(\delta_{12}^d = ?)$	-0.016*	-0.017	-0.007
		(0.008)	(0.017)	(0.015)
Number of Observations		218,046	50,419	49,214
Number of Firms		8,698	5,996	6,369
Adj. R-squared		0.17	0.17	0.17

† $|\beta_n^d|$ greater for constrained firms than unconstrained firms. ‡ φ^d greater for constrained firms than unconstrained firms.
§ $|\gamma_n^d|$ greater for constrained firms than unconstrained firms.