The Use of Trade Credit by Public and Private Firms: An Empirical Investigation

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Abstract

We show that the level of trade credit in private firms is one-quarter higher than that in public firms. The impact of being private on trade credit is robust, and is more pronounced in young, high-growth, and low-tangibility firms, consistent with the argument that firms with greater asymmetric information and credit constraints rely more on supplier financing. Both public and private firms seek to adjust toward optimal trade credit levels, although private firms experience slower adjustment. During the financial crisis of 2007–2009, public firms used slightly more trade credit as an alternative source of financing, while private firms were granted significantly less trade credit. (*JEL* G30, G31, G32, G01)

1. Introduction

Trade credit or accounts payable is a major component of working capital as nearly 40% of inventories and accounts receivable in US firms are financed with trade credit (Aktas, Croci, and Petmezas 2015). Just before the onset of the financial crisis of 2007, trade credit funded almost 90% of global merchandise trade, worth US$25 trillion (Klapper, Laeven, and Rajan 2012). Further, trade credit is the most important source of short-term financing for US firms (Petersen and Rajan 1997; Demirguc-Kunt and Maksimovic 2001), being three times larger than bank credit and fifteen times larger than commercial papers (Barrot 2014).[1] Trade credit also carries significant economic importance, acting as a substitute for bank credit during periods of monetary contractions or financial crises (e.g., Nilsen 2002; Choi and Kim 2005; Love, Preve, and Sarria-Allende 2007; Garcia-Appendini and Montriol-Garriga 2013).

Theory provides several motives for using trade credit (e.g., financing, transaction, and price discrimination), and explains why it is an important source of short-term financing. First, the supplier of trade credit has a cost advantage over specialized financial intermediaries because it knows more about, or has more control over, the buyer (Schwartz 1974; Emery 1984; Mian and Smith 1992; Petersen and Rajan 1997). According to this financing motive, financially unconstrained suppliers have a comparative advantage in extending trade credit to constrained buyers (Schwartz 1974), especially to those that face liquidity shocks that could endanger their survival (Cuñat 2007). In a similar vein, trade credit is beneficial for customers in distress because they are likely to be granted renegotiation concessions (Wilner 2000). The second, non-financial, motive for using trade credit is that it reduces transaction costs by separating the

[1] In a sample of non-financial US public firms for period 1987–1991, Rajan and Zingales (1995) find that accounts payable represent 15% of total assets, whereas short-term debt (debt in current liabilities) is 7.4% of total assets. Using our sample of both public and private firms, we document that the ratio of accounts payable to total assets for public (private) firms is 9.5% (16.1%), while the ratio of short-term debt to total assets is 6.4% (12.9%).

payment cycle from the delivery schedule, thus alleviating the problem with holding inventories of both money and goods (Ferris 1981). By allowing buyers to use a product before payment, trade credit also helps to reduce the costs associated with the verification of product quality (Smith 1987; Lee and Stowe 1993; Long, Maltiz, and Ravid 1993). Finally, trade credit can be used as a means for high-quality buyers to exercise their bargaining power (Giannetti, Burkart, and Ellingsen 2011), or for risky customers to obtain favorable price discrimination from suppliers (Brennan, Maksimovics, and Zechner 1988).

This paper investigates the use of trade credit by public and private firms. As reviewed above, previous studies show that trade credit is an important form of financing for firms with limited access to external capital markets (Petersen and Rajan 1997; Bias and Gollier, 1997; Berger and Udell 1998; Fisman and Love 2003). However, none of these studies has explored the impact of the listing status and, in particular, the importance of access to public equity markets on trade credit policy. This is a significant omission because a growing body of research has documented important differences in public and private firms' financial decisions, such as capital structure (Brav 2009), dividends (Michaely and Roberts 2012), cash holdings (Gao, Harford, and Li 2013), investments (Mortal and Reisel 2013; Asker, Farre-Mensa, and Ljungqvist 2015), and innovations (Gao, Hsu, and Li 2014; Acharya and Xu 2015). This literature also shows that public and private firms have fundamental differences in the degrees of asymmetric information, financial constraints, and creditworthiness, which are known factors affecting trade credit demands. This leads us to the following questions: Do private firms rely more or less on supplier financing than their public counterparts? Do firms have a target level of trade credit, and if so, does the listing status affect the speed with which they adjust toward this target? Did the credit shock associated with the recent financial crisis of 2007–2009 have

differential effects on trade credit in public and private firms? In the empirical work that follows, we seek to provide answers to these research questions.

We hypothesize that private firms will have a higher level of trade credit than their public counterparts. Compared to publicly listed companies, privately held firms have higher degrees of asymmetric information and financial constraints, and hence more limited access to alternative sources of financing (e.g., Brav 2009; Michaely and Roberts 2012; Gao, Harford, and Li 2013). Accordingly, private buyers will have a higher demand for trade credit than their public counterparts. Using a sample of US public and private firms collected from the S&P Capital IQ database for the period 1995–2012, we begin our analysis by documenting that the level of accounts payable in private firms is approximately one-quarter (23%) higher than that in their public counterparts. This result is both statistically and economically significant, and continues to hold for a sample of public firms and matched private firms.

We subject our finding to a battery of robustness checks. To address the potential sample selection and endogeneity concerns associated with the listing status of a company, we first analyze a transition sample of firms that were private and subsequently went public over the sample period. We observe a significant decline in these firms' reliance on trade credit post-listing, which is most pronounced in the first three years after the IPO. This finding is broadly in line with our baseline regression results. Next, we run a treatment regression that accounts for the endogeneity of the listing decision. Controlling for this endogeneity concern, our results continue to show a significantly higher level of trade credit in private firms than in their public counterparts. We then confirm our baseline regression results using propensity score matching. We also find that our findings are insensitive to including additional control variables, as well as using an extended sample and alternative measures of trade credit.

We next examine the impact of the listing status on the use of trade credit for subsamples of firms with different characteristics. We find that compared to publicly listed companies, privately held firms maintain a significantly higher level of trade credit, especially when they are younger, larger, and have more growth opportunities and fewer tangible assets. Young, high-growth, and low-tangibility firms often face high degrees of asymmetric information and credit constraints, which may explain their greater demand for trade credit. These findings are again consistent with our main argument based on asymmetric information and credit constraints.

In the second part of our analysis, we address the question about the optimum and dynamics of trade credit policy. We examine whether public and private firms attempt to adjust toward target levels of trade credit, and if so, whether the speed with which these firms adjust their trade credit varies according to the listing status. We find evidence to support the prediction that public firms move faster to target trade credit than their private counterparts, which is consistent with the notion that the former firms face lower adjustment costs than the latter. This result suggests that, although private firms rely more on trade credit than public firms, they may find it more difficult to adjust their trade credit and operate close to the optimal level.

In the final part of our empirical work, we investigate the impact of macroeconomic conditions on the use of trade credit by public and private firms. Following the literature, our expectation is that private firms resort to trade credit and rely more on this form of financing during a credit shock (Nilsen 2002). Examining the effects of the financial crisis of 2007–2009 on trade credit, we do not find evidence to support this argument. Our results show that private (public) firms were granted significantly less (slightly more) trade credit during the crisis. A possible explanation for these results is that the crisis affected not only financial lenders but also non-financial suppliers (Love, Preve, Sarria-Allende 2007), making it more difficult for private

firms to obtain supplier credit than their public counterparts. This finding shows how vulnerable private firms are during a credit crunch when the supply of trade credit, a potential substitute for bank credit, also dries up.

The main findings of our study contribute to the growing literature documenting differences in several important corporate financial policies between public and private firms (e.g., Brav 2009; Michaely and Roberts 2012; Gao, Harford, and Li 2013; Mortal and Reisel 2013; Gao, Hsu, and Li 2014; Acharya and Xu 2015; Asker, Farre-Mensa, and Ljungqvist 2015). To the best of our knowledge, we are the first to examine the use of trade credit by public and private US firms. Further, our research complements earlier studies of trade credit in small and medium-size firms (Petersen and Rajan 1997; Berger and Udell 1998; Giannetti, Burkart, and Ellingsen 2011). We note that those studies analyze data collected from the National Survey of Small Business Finance (NSSBF) in a single year (e.g., 1987, 1993 or 1998), and do not examine the impact of the listing status on the use of trade credit.

Our finding regarding the difference in the speed of adjustment between public and private firms contributes to the limited (non-US) literature studying the optimum and dynamics of various components of working capital, such as accounts receivable, accounts payable, and net trade cycle (García-Teruel and Martínez-Solano 2010a; García-Teruel and Martínez-Solano 2010b; Baños-Caballero, Garcia-Teruel, and Martinez-Solano 2014). It further adds to the recent evidence of differences in target adjustment behavior of public and private firms documented by research on other areas of corporate finance, such as capital structure (Brav 2009) and cash holdings (Gao, Harford, and Li 2013).

Finally, our analysis of the effect of the financial crisis on trade credit in public and private firms extends recent evidence for the former firm type documented by Love, Preve, and

Sarria-Allende (2007) and Garcia-Appendini and Montriol-Garriga (2013). Our study provides the first systematic evidence of the differential effects of the crisis on the use of trade credit by public and private firms, as well as relevant policy implications regarding the latter group.

The remainder of the paper is organized as follows. We review the literature in Section 2 and develop our hypotheses in Section 3. We discuss our data and the methodology in Section 4. We present our empirical results in Section 5 and conclude the paper in Section 6.

2. Related Literature

2.1. Trade credit: Theory and Evidence

One of the most important explanations for the use of trade credit is the financing motive (e.g., Schwartz 1974). In the presence of asymmetric information, the supplier of trade credit has a comparative advantage over traditional financial institutions in evaluating the buyer's creditworthiness and enforcing credit contracts (Cuñat 2007). Petersen and Rajan (1997) summarize this cost advantage in three dimensions: advantage in information acquisition, advantage in controlling the buyer, and advantage in salvaging value from the goods. First, compared to traditional lenders, the supplier of trade credit is able to obtain information about the buyer more quickly and at a lower cost thanks to the course of business between the two parties (Smith 1987; Biais and Gollier 1997; Burkart and Ellingsen 2004). Second, suppliers may control buyers by threatening them with cutting off their supplies; this threat is credible if there are limited sources of such supplies. Finally, trade creditors have a comparative advantage over traditional lenders in re-selling the goods in case of customer default thanks to their established network for selling within the industry (Mian and Smith 1992; Fabbri and Menichini 2010).

Among several theories of trade credit mentioned in the Introduction, the financing motive reviewed above is most relevant for research examining public and private firms because it is based on the assumption of asymmetric information between the supplier and the buyer, and is able to explain why trade credit is important for firms with credit constraints. To the extent that public and private buyers have varying degrees of informational asymmetries and financial constraints (e.g., Brav 2009; Schenone 2010; Michaely and Roberts 2012; Gao, Harford, and Li 2013), the incentives for suppliers to extend credit to these firms also vary.

There is an extensive body of empirical research examining various aspects related to the use of trade credit (e.g., Mian and Smith 1992; Long, Maltiz, and Ravid 1993; Ng, Smith, and Smith 1999; Danielson and Scott 2004; Choi and Kim 2005; Molina and Preve 2009; Klapper, Laeven, and Rajan 2012).[2] However, this literature focuses on either large, public firms or small and medium-size firms (hereafter SMEs). Using data collected from the NSSBF database, several recent studies have examined trade credit contracts among US SMEs (e.g., Petersen and Rajan 1997; Berger and Udell 1998; Giannetti, Burkart, and Ellingsen 2011). They generally find that supplier financing is important for firms that are financially constrained and have difficulty accessing other sources of external financing. While these studies provide useful insights about factors affecting not only the amount of trade credit but also details about credit contract terms, they have not examined the impact of the listing status of the borrowing firm on its demand for trade credit. This is an important gap in the trade credit literature because recent research examining public and private firms has revealed significant differences in the characteristics of

[2] There is a vast literature examining the use of trade credit in non-US SMEs. See, for example, Wilson and Summers (2002), Huyghebaert (2006), Niskanen and Niskanen (2006), Rodriguez-Rodriguez (2006), Cuñat (2007), García-Teruel and Martínez-Solano (2010c), and Baños-Caballero, Garcia-Teruel, and Martinez-Solano (2014). However, none of these studies examines the effect of the listing status.

these firm types, which have been shown by this literature to be important determinants of trade credit demands.[3]

2.2. Optimal trade credit and speed of adjustment to target trade credit

Emery (1984) shows how a firm derives its target trade credit by balancing the marginal benefits of trade credit against its marginal costs. From a buyer's perspective, trade credit brings about several benefits because it provides as an alternative source of financing for firms facing credit constraints (Schwartz 1974), liquidity shocks (Cuñat 2007), or financial distress (Wilner 2000), reduces transaction costs and provides a guarantee about product quality (Smith 1987), and allows certain buyers to obtain favorable price discrimination (Brennan, Maksimovics, and Zechner 1988). On the other hand, relying on supplier financing also has disadvantages because trade credit is generally a more expensive form of credit (Petersen and Rajan 1997), and there are opportunity costs associated with using trade credit due to a loss of discount for early payment (Ng, Smith, an Smith, 1999), or an increase in future cost of credit due to customer default (Nadiri 1969). Nadiri (1969) develops a theoretical model to study optimal trade credit policy by taking into account certain benefits and costs of trade credit. He argues that the observed level of trade credit may deviate from the optimum due to firms' inaccurate estimates of sales, purchases, and the opportunity costs of trade credit, as well as disequilibrium in other assets such as inventories. However, firms should attempt to close out any deviation from the optimal level by making adjustment in their trade credit over time.

[3] Atanasova (2007) examines the impact of credit constraints on the use trade credit for a small sample of UK public and private firms. Her analysis, however, uses a regime-switching approach to categorize firms into constrained vs. unconstrained, and thus does not focus on examining the effect of the listing status on trade credit. Using a sample of European public and private firms, Anagnostopoulou (2012) shows that public firms have a higher cash conversion cycle than their private counterparts. However, the author uses a broad measure of working capital and does not examine its components including trade credit.

Based on Nadiri's (1969) adjustment framework, a few papers have examined the dynamics of trade credit and related components of working capital, such as accounts receivable (García-Teruel and Martínez-Solano 2010a), or net trade cycle (Baños-Caballero, Garcia-Teruel, and Martinez-Solano 2014). Particularly, García-Teruel and Martínez-Solano (2010b) investigate the speed of adjustment toward optimal accounts payable using a sample of UK SMEs; they find that these firms adjust quickly to their target trade credit with a speed of 77%. However, all these studies use non-US data, and none of them has examined potential different adjustment behavior of public and private firms. This is an important omission because recent research on public and private firms has investigated target adjustment behavior of other financial policies. In the capital structure literature, Brav (2009) shows that public firms adjust toward target leverage more quickly than their private counterparts. However, in recent research on cash holdings, Gao, Harford, and Li (2013) find that public firms are slower than private firms in moving toward their target cash balances.

2.3. The effect of macroeconomic conditions on trade credit

The use of trade credit is affected by macroeconomic conditions (Schwartz 1974; Smith 1987), and this effect varies according to the creditworthiness of the firm; see extensive reviews by Minshkin (1995) and Mateut (2005). According to prior research, trade credit can act as a substitute for other sources of external financing such as bank loans, especially during periods of monetary contractions (Bias and Gollier 1997; Petersen and Rajan 1997; Nilsen 2002; Kim and Choi 2005). This finding is based on Meltzer's (1960) redistribution view that large, liquid firms with better access to capital markets will have incentives to redistribute the credit received to financially less secure firms via trade credit. In Cuñat's (2007) theoretical model, suppliers are willing to extend trade credit to buyers faced with a liquidity shock, even if traditional lenders

refuse to do so. Empirically, Nilsen (2002) finds that small firms and even large firms without a credit standing resort to trade credit at times of tight monetary policy.

Recent research has examined the effects of crises on the use of trade credit. Love, Preve, and Sarria-Allende (2007) examine trade credit during the Asian and Mexican currency crises. They find that the amount of trade credit provided and received increased at the peak of those crises, although it contracted post-crisis due to the shrinking of both bank and supplier credit. Most recently, Garcia-Appendini and Montriol-Garriga (2013) show that during the 2007–2008 financial crisis, liquid suppliers extended more trade credit to support their customers and consequently experienced better performance. On the demand side, trade credit taken, especially by constrained firms, also increased during the credit crunch.

3. Hypotheses

In this section, we develop three hypotheses corresponding to the research questions set out in our Introduction. To begin with, we expect to find a significant difference in the level of trade credit between public and private firms because there are well-documented differences between these two firm types. First, due to information disclosure requirements, publicly listed firms are more transparent and have lower degrees of asymmetric information than their privately held counterparts (Brav 2009; Schenone 2010). Since suppliers can acquire information about opaque firms more quickly and control them with supplier financing more effectively than formal lenders (Smith 1987; Biais and Gollier 1997; Burkart and Ellingsen 2004; Petersen and Rajan 1997), they have a financing motive to extend trade credit to private firms.

Second, private firms are more financially constrained than public firms because they have no access to public equity markets. They also face a higher cost of debt (Brav 2009; Gao, Harford, Li 2013), and have weaker bargaining power with banks (Saunders and Steffen 2011).

Hence, private firms have greater incentives to use trade credit as a substitute for other forms of external financing. On the supply side, the demand for trade credit by private firms is likely to be accommodated by unconstrained suppliers, at least during normal times, because trade creditors have incentives to exploit a cost advantage over specialized financial institutions by redistributing their credit (Petersen and Rajan 1997).

Third, prior evidence suggests that private firms have a lower credit quality and are riskier than public firms (Pagano, Panetta, and Zingales 1998; Scherr and Hulburt 2001). While independent lenders are reluctant to provide credit to firms with a high failure rate, suppliers are willing to do so (Wilner 2000). As a result, private firms have strong incentives to rely on trade credit as one of the most important source of short-term financing. Also note that trade credit is more flexible than bank loans because it is easier to renegotiate due to its revolving nature, and is less costly to delay repayments (Wilner 2000; Cuñat 2007). In the event of a customer default, trade creditors often provide more concessions to maintain their product market relationship (Wilner 2000). Petersen and Rajan (1997) argue that the implicit equity stake of suppliers in buyers consists of the present value of current and future sales. Hence, unlike banks, suppliers are concerned about the continuation of their customers and tend to be lenient toward cases of financial distress (Huyghebaert, Van De Gucht, and Van Hulle 2007). These benefits are valuable for private firms with a lower credit quality and a higher risk than their public counterparts. Based on the above arguments, we develop the following hypothesis:

H1: Private firms have a higher level of trade credit than public firms.

According to pioneering research by Nadiri (1969) and Emery (1984), we predict that a firm will have an optimal level of trade credit that balances the benefits and costs of trade credit and maximizes the firm's value. Further, we argue that the speed with which firms adjust toward

this optimum depends on the costs of adjusting trade credit. We expect public firms to have lower adjustment costs than private firms because the former firms may have greater bargaining power, allowing them to renegotiate with suppliers, adjust the amount of trade credit taken, and amend credit contract terms more easily (Klapper, Laeven, and Rajan 2012). Public firms also face a lower cost of capital (Campello et al. 2011; Saunders and Steffen 2011), as a result of having greater transparency (Schenone 2010), greater liquidity (Pagano, Panetta, and Zingales 1998), and greater access to external sources of liquidity (Faulkender and Petersen 2006; Lins, Servaes, and Tufano 2010). Having a lower cost of capital allows public firms to adjust their trade credit more easily and quickly by switching to other forms of credit. In sum, this argument suggests that private firms should have a lower speed of adjustment than public firms.

However, one could argue that public firms have less incentive to operate at target trade credit than private firms, especially when the costs of deviating from such target are not material to them. For instance, in theory, firms maintaining more trade credit than the optimal level face the expected costs of default and higher costs of future credit due to the deterioration in their credit reputation. However, to the extent that public firms are less prone to bankruptcy and have better bargaining power than private firms, they are less concerned about the possibility of default as well as the costs of deviating from target trade credit. As a result, the incentive for public firms to revert toward the optimal trade credit may be weaker than that for private firms. Overall, given the above conflicting arguments, we develop the following alternative hypotheses:

H2a: Private firms adjust toward target trade credit more slowly than public firms.

H2b: Private firms adjust toward target trade credit more quickly than public firms.

According to prior research examining the impact of monetary contractions and financial crises on the use of trade credit, we argue that both public and private firms will have incentives

to substitute bank financing for supplier financing during a credit crunch, when the supply of bank credit dries up. Our argument is based on the redistribution view of trade credit provision (Meltzer 1960; Petersen and Rajan 1997) and empirical evidence showing public firms increased their trade credit at the peak of the Asian and Mexican currency crises (e.g., Love, Preve, and Sarria-Allende 2007), as well as the recent financial crisis (Garcia-Appendini and Montriol-Garriga 2013). Moreover, we also predict that between public and private firms, the latter firms will have a greater demand for trade credit during a credit shock. This is because private firms are more constrained than public firms, and have even more limited access to alternative sources of financing. Our prediction is also motivated by prior evidence on the differential impacts of macroeconomic uncertainty on firms with varying degrees of constraints (Korajczyk and Levy 2003), and in particular, results regarding the effect of monetary tightening on the use of trade credit by constrained firms (Nilsen 2002; Atanasova 2007; Garcia-Appendini and Montriol-Garriga 2013). In sum, our arguments enable us to develop our final hypothesis:

H3: Private firms experience a greater increase in the level of trade credit than public firms during a financial crisis.

4. Data and Methodology

4.1. Sample description

We collect our data from the S&P Capital IQ database for the period 1995–2012. S&P Capital IQ provides data on both US public and private firms;[4] however, its coverage of trade

[4] S&P Capital IQ provides data on private firms that file Forms 10-K (annual reports), 10-Q (quarterly reports), or S-1 (securities registration) with the Securities Exchange Commission (SEC). According to the SEC regulations, firms with total assets of $10 million or above, and with 500 or more shareholders are required to file 10-K and 10-Q reports, while firms with public debt are required to file S-1 Form. S&P Capital IQ confirms that they also cover other private firms from the third-party Private Company Financials provider, which receives data by directly

credit includes more private than public firms. Colla, Ippolito, and Li (2013) compare the quality of data for public firms provided by Compustat and that provided by S&P Capital IQ. Examining several corporate variables such as leverage, size, profitability, cash holdings, tangibility, and asset maturity, they conclude that the quality of S&P Capital IQ data is comparable to that of Compustat data. In Table A.1, we show that Colla, Ippolito, and Li's (2013) finding can be extended to data on trade credit. Specifically, we find that the summary statistics of the trade credit variable for public firms collected from S&P Capital IQ and from Compustat are comparable. This suggests that our sample of public firms from S&P Capital IQ is a representative sample of all public firms in Compustat.

Following Gao, Harford, and Li (2013), we exclude financials and utilities, IPO firms, firms that went private during our sample period, and firms with a cash flow to total assets ratio of less than -50%. Next, we remove observations with missing variables, negative equity, and negative total assets. Finally, we winsorize the variables at the 1st and 99th percentiles to alleviate the impact of outliners. Our sample consists of 27,300 private firms with 70,011 firm-year observations and 3,340 public firms with 33,766 firm-year observations.[5]

4.2. Empirical models

To examine differences in the trade credit ratios of public and private firms (*Hypothesis 1*), we estimate the following model:

$$TC_{it} = \beta_0 + \beta_1 Public_{it} + \boldsymbol{\theta}' \boldsymbol{X}_{it} + \varepsilon_{it}, \quad (1)$$

contacting the company, from CPAs, from courts and recording offices regarding suits, liens, judgments and bankruptcy filings, and from top news providers. The use of various sources for data on private firms provides more coverage and alleviates the sample selection concern.

[5] Consistent with Asker, Farre-Mensa, and Ljungqvist (2015) and Gao, Lemmon, and Li (2012), our sample includes more firm-year observations for private firms than for public firms.

where the dependent variable, trade credit (TC_{it}), is measured as the ratio of accounts payable to total assets (Petersen and Rajan 1997; Fisman and Love 2003; Giannetti, Burkart, and Ellingsen 2011). *Public* is a dummy variable that takes the value of 1 for public firms, and 0 for private firms. Following prior research (Petersen and Rajan 1997; Love, Preve, and Sarria-Allende 2007; Garcia-Appendini and Montoriol-Garriga 2013), our control variables ($\boldsymbol{X}_{it}$) include ln(1+age), ln(1+age)^2, cash flow, cash holdings, current assets, negative growth, positive growth, short-term debt, and firm size. We provide a detailed discussion of these variables in Section 4.3. We note that our model is similar that used by Gao, Harford, and Li (2013) to investigate differences in the cash holdings of public and private firms. Since we hypothesize that private firms rely more on trade credit than their public counterparts, we expect β_1 to be negative.

Next, to test *Hypotheses 2a* and *2b*, we compare how quickly public and private firms adjust toward their target trade credit levels. We do so by estimating the following partial adjustment model for public and private firms separately:

$$\Delta TC_{it} = \beta_0 + \delta(TC_{it}^{*} - TC_{it-1}) + \varepsilon_{it} . \quad (2)$$

The dependent variable, ΔTC_{it}, is the change in trade credit from year $t-1$ to t. TC_{it-1} is the lagged value of trade credit. TC_{it}^{*} is the target trade credit ratio, which is estimated from a regression of trade credit on the control variables listed above, separately for public and private firms, as follows:

$$TC_{it} = \beta_0 + \boldsymbol{\gamma}'\boldsymbol{X}_{it} + \varepsilon_{it}. \quad (3)$$

Our approach allows for the possibility that public and private firms may maintain heterogeneous trade credit targets, which is in line with our argument that these firms have different trade credit demands. The coefficient of interest δ measures the speed of adjustment toward the target level of trade credit; it takes a value from zero to one. If the firm adjusts its trade credit immediately,

the speed of adjustment will be equal to one. However, the speed of adjustment will be equal to zero if the adjustment costs are so high that the firm is unable to adjust to its target trade credit. We compare the speed of adjustment between public and private firms by testing whether the difference in the estimates of δ is statistically significant using the Chow test.

Finally, to examine the differential effects of credit conditions on public and private firms' use of trade credit (*Hypothesis 3*), we estimate following model:

$$TC_{it} = \beta_0 + \beta_1 Crisis_{it} + \beta_2 Public_{it} + \beta_3 Crisis_{it} \times Public_{it} + \boldsymbol{\theta}' \boldsymbol{X}_{it} + \varepsilon_{it}. \quad (4)$$

Model (4) extends model (1) in that it includes dummy variables to account for the effects of a supply credit shock, proxied by the recent financial crisis of 2007–2009.[6] $Crisis$ is a dummy variable that takes the value of 1 in the years 2007–2009, and 0 otherwise. Our chosen period consists of the first-stage of the crisis from July 2007 to June 2008 and the second-stage following the bankruptcy of Lehman Brothers in September 2008 until the fourth quarter of 2009. To avoid cofounding effects due to other periods of macroeconomic fluctuations before 2003, we estimate the model for the period 2004–2009. $Crisis * Public$ is an interaction term between the *Public* and *Crisis* dummy variables. The effects of the crisis on private firms and public firms are captured by β_1 and $\beta_1 + \beta_3$, respectively. We expect both public and private firms to rely more on trade credit during the crisis with private firms' use of trade credit increasing the most, i.e., $\beta_1 > 0$ and $\beta_3 > 0$.

4.3. Control variables

Consistent with the literature, the control variables in models (1), (3), and (4) include ln(1+age), ln(1+age)^2, cash flow, cash holdings, current assets, negative growth, positive growth, short-term debt, and size. A firm's demand for trade credit is affected by its

[6] We obtain qualitatively similar results when we restrict the crisis period to be 2007–2008.

creditworthiness, which can be proxied by firm size and age (Petersen and Rajan 1997). Large and old firms may use less trade credit because they have access other sources of financing. However, from a supplier's perspective, these firms are more likely to be granted trade credit with favorable terms due to their better credit quality (Klapper, Laeven, and Rajan 2012). These alternative arguments suggest that the direction of the relationship between trade credit and size (age) should be resolved empirically. Consistent with Scherr and Hulburt (2001), we measure size as the natural logarithm of total sales in 2012 dollar price. Following Petersen and Rajan (1997), age is the number of years from incorporation, and is calculated as ln(1+age). We also include $\ln(1+age)^2$ to account for the conjecture that early years are more important for building a firm's reputation than later years; we expect it to be negatively related to trade credit.

The pecking order theory shows that, in the presence of asymmetric information, firms prefer to use internal funds to external financing (Myers and Majluf 1984). Since trade credit is an external source of financing, this theory predicts that firms will resort to trade credit after having exhausted internal resources. We thus expect trade credit to have a negative relationship with internal funds. As in Petersen and Rajan (1997), we measure the internal sources of financing using cash flow, calculated as the ratio of net profits plus depreciation to total assets.

Morris (1976) argues that firms should match the maturities of their assets and liabilities in order to ensure that cash inflows generated from the assets can always cover the cash outflows to service the liabilities. Myers (1977) further shows how firms can reduce underinvestment incentives and the agency costs of debt by matching the maturity structure of their debt to the life of their assets. Based on these arguments, we expect trade credit and current assets to have a positive relationship; we calculate currents assets as current assets minus cash, divided by total assets (Petersen and Rajan 1997). In addition, we follow Love, Preve, and Sarria-Allende (2007),

and Garcia-Appendini and Montriol-Garriga (2013) and include cash holdings as an additional control variable. We expect cash holdings to have a positive effect on trade credit; we measure cash holdings as the ratio of cash and cash equivalents to total assets.

High-growth firms tend to be more constrained and, as a result, rely more on trade credit (Cuñat 2007). Alternatively, firms with high growth, especially in sales, have a greater demand for trade credit to finance the new investments in current assets. Following previous studies of private firms (Scherr and Hulburt 2001; Brav 2009), we measure growth using sales growth. Consistent with Petersen and Rajan (1997), we differentiate between positive and negative growth; positive (negative) growth is defined as sales growth multiplied by a positive growth dummy variable that takes the value of 1 if sales growth is positive (negative), and 0 otherwise. We expect trade credit to be positively (negatively) related to positive (negative) growth.

Prior research shows how trade credit acts as a substitute for other sources of short-term financing (Petersen and Rajan 1997; Nilsen 2002). Hence, to model the demand for trade credit, we include short-term debt to control for alternative forms of credit. We expect short-term debt to have a negative relation with trade credit to reflect the substitution effect. Following previous research on private firms (Scherr and Hulburt 2001), we measure short-term debt as the ratio of short-term borrowings plus the current portion of long-term debt to total assets.

5. Empirical Results

In this section, we first report the univariate and multivariate results regarding the difference in trade credit policies between public and private firms. Next, we present regression results conditional on firm characteristics. We then examine whether and how quickly public and private firms adjust toward their target trade credit. Finally, we provide evidence on the effects of macroeconomic conditions on the use of trade credit by public and private firms.

5.1. Difference in trade credit between public and private firms

5.1.1. Sample overview and Univariate analysis

Panel A of Table 1 reports the summary statistics for the full sample of public and private firms. The mean trade credit ratio is 13.6% (median of 8.9%), which is much higher than the figure for short-term debt (mean of 3% and median of 0%). Current assets represent more than a half of total assets (50.7%), while cash holdings is 13.4% of total assets. The average age of the sample firms is about 40 years (median of 31 years).

Panel B of Table 1 presents the results from our univariate analysis. Public firms have a mean trade credit of 8.9% (median of 6.6%), which is about half of the figure for private firms (mean of 15.8% and median of 11.1%).[7] The difference in the trade credit ratios of the two groups is 6.9% (median of 4.5%), and is significant as confirmed by our statistical tests. This provides the first evidence to support *Hypothesis 1* that private firms rely more on trade credit than their public counterparts. In unreported analysis, we also observe significant differences in other characteristics of public and private firms, consistent with the literature. For example, public firms are larger and more mature, as well as have less short-term debt, less cash flow, and more cash holdings than their private counterparts (see Brav 2009; Gao, Harford, and Li 2013; Gao, Hsu, and Li 2014; Asker, Farre-Mensa, and Ljungqvist 2015).

[Insert Table 1 here]

Table 2 shows how trade credit varies across the 12–Fama French industries. Several studies argue that the use of trade credit is uniform within an industry but varies across industries (Smith 1987; Ng, Smith, and Smith 1999; Fisman and Love 2003). Panel A shows that firms relying most on trade credit operate in the retail and wholesale industry, with a mean trade credit

[7] The ratio of account payables to total liabilities 21.7% (33.8%) in public (private) firms.

ratio of 18.8% (median of 14.7%). This finding is similar to García-Teruel and Martínez-Solano's (2010c) earlier evidence for European firms. Firms with the lowest level of trade credit (mean of 6.5% and median of 4.1%) are in the health sector (including healthcare, medical equipment, and drugs). This finding is consistent with Fisman and Love (2003), who argue that trade credit is unpopular for drugs companies because it is difficult for suppliers to resell these specific products in the event of a customer default.

Panel B of Table 2 presents the summary statistics of trade credit across industries. The health sector still exhibits the lowest level of trade credit, and also the smallest difference in the level of trade credit between public and private firms. The largest difference in public and private firms' trade credit ratios is observed in the energy industry (difference in mean of 8% and in median of 3.6%). The difference in the level of trade credit between public and private firms in business equipment is almost as high (difference in mean of 7.5% and in median of 4.2%). The difference in the trade credit ratios of public and private firms in other industries ranges between 3.2% and 4.6%. More importantly, the test statistics for differences in mean and median are all statistically significant. Overall, we find that public and private firms maintain significantly different levels of trade credit and this finding holds across industries.

[Insert Table 2 here]

5.1.2. Multivariate analysis – Baseline regression results

Table 3 reports the regression results for model (1), in which we investigate the difference in the use of trade credit by public and private firms. In the first two columns, we simply regress trade credit on the *Public* dummy without controlling for firm-specific characteristics. The results in Column (1) show that the difference in trade credit between public

and private firms (6.9%) is significantly negative.[8] In Column (2), we include industry and year effects to account for unobserved industry-level heterogeneity and common time trends.[9] We find that the difference in public and private firms' trade credit decreases to 4.4%, but remains statistically significant. In Column (3), we include the full set of control variables. In Column (4) we further control for industry and year effects.[10] The results show that, even after controlling for firm-specific variables, the *Public* dummy is still significantly negative. The difference in the level of trade credit in public and private firms varies between 5.5% and 3.6%, and is economically significant. Using the *Public* dummy coefficient estimate in the baseline model in Column (4), the difference in public and private firms' trade credit is 23% relative to the mean trade credit ratio of private firms (15.9%). These finding strongly supports *Hypothesis 1* that privately held firms have a higher level of trade credit than their publicly listed counterparts. In Column (5), we estimate the baseline model again using a matched sample of public and private firms. In particular, we employ a *one*-to-*n* matching technique, where we match, with replacement, each public firm-year observation to any private firm-year observation in the same industry and year, and of similar size (allowing for a deviation of 30%). The results show that the difference in the trade credit ratios of the public and matched private firms (6.6%) remains statistically significant, and even becomes economically stronger.

[Insert Table 3 here]

[8] In our regressions, we report heteroskedasticity-consistent (robust) standard errors. However, the results remain qualitatively the same if we estimate robust standard errors clustered at the firm level.

[9] In unreported analysis, we also examine the evolution of trade credit for both public and private firms (in Figure 1 we present different graphical evidence of how trade credit evolves around the IPO). However, unlike recent evidence of the secular increase in cash holdings (Bates, Kahle, and Stulz 2009) and in short-maturity debt (Custódio, Ferreira, and Laureano 2013), we observe no clear pattern in trade credit over time.

[10] Since the firms in our sample remain public or private throughout the sample period, we do not include firm fixed effects because they are highly correlated with the *Public* dummy variable. Our approach is consistent with Gao, Lemmon, and Li (2012), Gao, Harford, and Li (2013), and Asker, Farre-Mensa, and Ljungqvist (2015).

The results regarding the control variables are broadly consistent with the literature. ln(1+age) has a significantly negative coefficient, indicating that young firms use more trade credit. However, we find that $\ln(1+age)^2$ is insignificant. As expected, firms with high cash flow rely less on trade credit, consistent with the evidence in Petersen and Rajan (1997). The results also show that current assets and trade credit are positively related. Similarly, cash holdings have a positive impact on trade credit, in line with Garcia-Appendini and Montriol-Garriga (2013). Both of these results support the principle of matching the maturities of assets and liabilities. Positive growth has a positive effect on trade credit, consistent with our prediction and Petersen and Rajan (1997). However, negative growth is also significantly positive, suggesting that firms with negative growth options are provided with more trade credit. A possible explanation for this finding is that firms that have few growth opportunities but are large and less constrained can borrow more (Petersen and Rajan 1997). Firms with high short-term debt have less trade credit, which is in line with the substitution effect. Finally, large firms are granted more trade credit, supporting the notion that suppliers are willing to grant more favorable and longer credit terms to large firms because these firms have greater market power (Klapper, Laeven, and Rajan 2012).

5.1.3. Dealing with sample selection and endogeneity

Transition sample

To deal with the sample selection concern about firms self-selecting themselves to go public, we use a transition sample of the same IPO firms that went public during the sample period. Using this transition sample mitigates the sample selection bias by controlling for the selection on the time-invariant unobservable firm characteristics; see also Michaely and Roberts (2012) and Gao, Harford, and Li (2013) for a similar approach.

Figure 1 demonstrates graphically the evolution of the trade credit policy of IPO firms around the IPO event, specifically from the IPO–4 year to the IPO+5 year. The peak mean trade credit is in the IPO–2 and IPO–1 years, both at 9.1% (median of 5.1% and 5.8%, respectively). In the IPO year, there is a 1.4% decline in the mean trade credit, which remains stable at this relatively low level, compared to the pre-IPO level of about 7%, until IPO+5. Overall, this graphical evidence is in line with the multivariate analysis above.

[Insert Figure 1 here]

Next, we perform a regression analysis to investigate the change in trade credit post-IPO using this transition sample. We estimate the permanent effect of the listing decision on trade credit post-IPO as well as the temporary effects in the first few years following the IPO. Our estimated models are specified as follows:

$$TC_{it} = \beta_0 + \beta_1 D_{Post\_IPO} + \boldsymbol{\theta}' \boldsymbol{X}_{it} + \varepsilon_{it}. \quad (5)$$

$$TC_{it} = \beta_0 + \beta_1 D_{IPO} + \beta_2 D_{IPO+1} + \beta_3 D_{IPO+2} + \beta_4 D_{IPO+3} + \boldsymbol{\theta}' \boldsymbol{X}_{it} + \varepsilon_{it}. \quad (6)$$

In both models, the dependent variable is trade credit. In model (5), D_{Post_IPO} is a dummy variable that takes the value 1 in the IPO year and the years after the IPO, and 0 otherwise. In model (6), D_{IPO} is a dummy variable that takes the value 1 in the IPO year, and 0 otherwise; D_{IPO+i} with i=1..3 is a dummy variable that takes the value 1 in the IPO+i year, and 0 otherwise. $\boldsymbol{X}_{it}$ is a vector of the control variables as discussed above.

Table 4 presents the regression results for the transition sample of 1,282 IPO firms that did an IPO during our sample period. Columns (1)–(3) report the results for model (5), without and with industry, year, and firm effects, respectively.[11] The D_{Post_IPO} dummy is significantly negative, which lends support to the notion that as the firm goes public its reliance on trade credit

[11] Using year dummies also accounts for time trends in the market conditions that may affect the IPO timing.

decreases. The magnitude of the decline in trade credit is between 2.2% and 1.1%, which remains economically significant. In Columns (4)–(6), we report the results regarding the temporary effects of the IPO decision on trade credit, without and with industry, year, and firm effects, respectively. The results suggest that IPO firms rely significantly less on trade credit in the first few years post-IPO. Specifically, compared to the pre-IPO level, firms reduce their trade credit level by 1%–2.1% between IPO and IPO+3. Overall, the results for the transition sample suggest that IPO firms maintain a significantly lower level of trade credit after going public, which lends further support for *Hypothesis 1*.

[Insert Table 4 here]

Treatment regression

Going public can be an endogenous decision because it may be affected by unobserved firm characteristics that are also related to trade credit. To deal with this potential endogeneity problem, we run a treatment regression that involves estimating the following models in two stages (Gao, Harford, and Li 2013):

$$Public = \gamma_0 + \boldsymbol{\gamma}'\boldsymbol{Z}_{it} + \omega_{it}; \qquad First-stage\ regression \tag{6}$$

$$TC_{it} = \beta_0 + \beta_1 Public_{it} + \boldsymbol{\theta}'\boldsymbol{X}_{it} + \varepsilon_{it}\,. \quad Second-stage\ regression \tag{7}$$

In the first stage, we estimate a probit model (6), which captures the decision to go public. The second stage involves estimating model (7), in which trade credit is regressed on the fitted values of the *Public* dummy variable, estimated from the first stage, and the control variables. Following Gao, Harford, and Li (2013), we use industry-level underwriter concentration as an instrument for the *Public* dummy. This instrumental variable affects the costs of doing an IPO and subsequently the listing decision (Liu and Ritter 2011), but is not related to other corporate

decisions such as trade credit.[12] For identification purposes, we include all the exogenous regressors in the first-stage regression along with this instrument.

We tabulate the treatment regression results obtained using the maximum likelihood estimator in Table 5, with the first-stage regression results in Column (1) and the second-stage regression results in Column (2). In Column (1), the underwriter concentration variable is statistically significant and has the expected negative sign. This suggests that the higher the underwriter concentration, the lower the costs of doing an IPO, and the more likely firms are listed. We also note that the diagnostic test statistic is significant, supporting the notion that the going public decision should be treated as endogenous and validating our treatment regression approach. In Column (2), the results regarding the *Public* dummy suggest that public firms maintain a significantly lower level of trade credit than private firms. The difference (7%) in public and private firms' trade credit is economically stronger than the baseline result. This suggests that, failing to control for the endogeneity concern may underestimate the impact of the listing status on the use of trade credit. Overall, our finding that private firms use a higher level of trade credit than public firms is robust to accounting for endogeneity.

[Insert Table 5 here]

Propensity score matching

Table 6 presents the results of the propensity score matching analysis, which controls for selection based on observable differences between public and private firms. Since we have more observations for private firms, we consider public firms as the treated group and private firms as the control group. Using the propensity score matching technique, we implement *one*-to-*one*

[12] Industry-level underwriter concentration is defined as the number of IPOs underwritten by the top five underwriters divided by the number of IPOs in that industry. The top five underwriters are determined by the number of IPOs underwritten in the last five years (Liu and Ritter 2011).

matching to the nearest neighborhood, without replacement (Gao, Harford, and Li 2013).[13] Specifically, we match each public firm-year observation to a private firm-year observation using the propensity score of being public from a probit regression based on certain firm characteristics. In the probit model, we use two matching specifications to capture the status of being a privately held or publicly listed firm. In Specification 1, the matching is based on firm size, industry, and year effects. In Specification 2, the matching is based on all the control variables used in our baseline regression, including ln(1+age), $\ln(1+age)^2$, cash holdings, cash flow, current assets, negative growth, positive growth, short-term debt, size, and industry and year effects. Panel A presents the pairwise differences in the mean trade credit in the propensity score-matched sample together with the bootstrapped standard errors based on 50 replications. We find that in Specification 1(2), there is a statistical and economic difference in the trade credit of the propensity score-matched public and private firms, i.e., the level of trade credit in private firms is 7.9% (6.1%) higher than that in public firms. In Panel B, we re-estimate the baseline regression model using the propensity score-matched sample. The results confirm our earlier baseline results that private firms have a higher level of trade credit (by 3.6%–3.7%).

[Insert Table 6 here]

5.1.4. Additional robustness checks

Table 7 reports several robustness checks in which we include additional control variables, consider an extended sample, and use an alternative measure of the dependent variable. In the first two columns, we include two additional control variables, namely the cost of

[13] We obtain qualitatively similar results when we use propensity score matching without replacement, and a caliper of 30% standard deviation of the propensity score.

external finance and the annual rate of GDP growth.[14] Specifically, we control for the cost of external finance in Column (1), consistent with previous studies examining small firms (e.g., Rodriguez-Rodriguez 2006; García-Teruel and Martínez-Solano, 2010b). The higher the cost of external finance, the higher is the demand for trade credit. We measure the cost of external finance (Fcost) as interest expense divided by total debt minus trade credit. The results show that this variable is positively related with trade credit as expected.

In Column (2), we include the annual rate of GDP growth to proxy for macroeconomic conditions (e.g., Niskanen and Niskanen 2006; García-Teruel and Martínez-Solano 2010b). During adverse conditions, trade credit is an alternative source of funds and, thus, its demand is expected to increase. At the same time, during favorable macroeconomic conditions, there are more investment opportunities, which require more funding including trade credit. GDP data comes from the Federal Reserve Bank of St Louis website. Consistent with the former argument, we find a positive relation between the GDP growth rate and trade credit.

In Column (3), we follow prior research (e.g., Petersen and Rajan 1997) and analyze a larger sample of firms that includes utilities. In Column (4), we follow Love, Preve, and Sarria-Allende (2007) and Garcia-Appendini and Montriol-Garriga (2013) and use an alternative measure of trade credit in which we normalize accounts payable by the cost of goods sold, instead of total assets as in our main analysis.

Overall, the above robustness checks show that our main findings continue to hold. Indeed, there is strong and robust evidence that private firms rely more on trade credit than their public counterparts. The difference in their trade credit varies between 2.2% and 3.8%, which is similar in magnitude to the results for the baseline model (3.6%).

[14] In unreported results, we include inventory as an additional control variable and obtain qualitatively similar results.

[Insert Table 7 here]

5.2. Results conditional on firm-specific characteristics

In this section, we examine whether the difference in public and private firms' trade credit varies according to certain firm-specific characteristics that proxy for the degrees of asymmetric information and credit constraints, and are related to the financing motive of supplier financing. We consider subsamples of firms according to their age, growth opportunities, tangibility, and size, respectively (see notes to Table 3 for the definitions of these variables). Table 8 presents the results for models using dummy variables and their interaction terms with the *Public* dummy. The *Mature* dummy variable takes the value of 1 for firm-year observations with above the median age, and 0 otherwise. The *High growth* dummy variable takes the value of 1 for observations with above the median growth, and 0 otherwise. The *High tangibility* dummy variable takes the value of 1 for observations with above the median tangibility, and 0 otherwise. The *Large* dummy variable takes the value of 1 for observations with above the median size, and 0 otherwise. We expect young, high-growth, low-tangibility, and small firms to have greater information problems and less access to external capital markets.

The results in Column (1) show that the *Public* dummy variable remains statistically negative, in line with our earlier findings. While *Mature* is significantly negative, its interaction term with *Public* dummy is significantly positive. Consistent with our conjecture, this finding suggests that young firms maintain a higher level of trade credit than their mature counterparts. Further, within the group of young firms, those that are private rely more on trade credit than those that are public. The results in Columns (2)–(3) show that firms with high growth and low tangibility use more trade credit than those with the opposite characteristics. Further, within these groups of firms, those that are private have a higher level of trade credit. Private firms that

are young, and have high growth options, or limited tangible assets face higher degrees of asymmetric information and financial constraints, which may explain why these firms have greater demands for trade credit. This finding is consistent with previous evidence documenting the impact of asymmetric information and credit constraints on the use of trade credit (Nilsen 2002; Fisman and Love 2003; Cuñat 2007). In Column (4), we find that large firms use more trade credit than small firms, which is consistent with some evidence in the literature (Petersen and Rajan 1997; Klapper, Laeven, and Rajan 2012). This finding also supports a supply-side explanation that large firms are granted more trade credit due to their better credit quality and greater bargaining power. However, within the group of public firms, those that are larger rely less on trade credit. Considering that large firms are less concerned about information problems than small firms, this latter finding is in line with the argument based on asymmetric information, and the results in Columns (1)–(3). Overall, our results suggest the difference in public and private firms' use of trade credit varies with the levels of informational asymmetries and credit constraints facing these firms.

[Insert Table 8 here]

5.3. Speed of adjustment to target trade credit

In this section, we examine whether public and private firms adjust toward target trade credit levels and whether they do so with different adjustment speeds. We first estimate the target level of trade credit separately for public and private firms and tabulate the results in Table A.2 of the Appendix. Our approach accounts for a difference in the target trade credit level between

public and private firms; indeed, we find evidence of significantly different coefficients on the factors affecting those target levels.[15]

Panel A of Table 9 presents the regression results for the partial adjustment model of trade credit. The results indicate that both public and private firms adjust toward their target levels of trade credit at moderate rates. This result is consistent with the argument that firms have optimal trade credit (Nadiri 1969) and that they seek to adjust toward this target. Empirically, our estimated speeds of adjustment are statistically significant, but much lower in magnitude than the speed of adjustment estimated using a sample of UK SMEs in previous research (García-Teruel and Martínez-Solano 2010b). More importantly, we find that private firms adjust at a slower rate (23%) than public firms (29%); the Chow test also confirms that the difference in these adjustment speed estimates is statistically significant. This finding supports *Hypothesis 2a* that private firms have a slower adjustment speed due to facing higher adjustment costs. The finding that private firms adjust their trade credit more slowly than their public counterparts is in line with the earlier evidence on leverage adjustment (Brav 2009).

In Panels B and C, we investigate whether the difference in the speed of adjustment between public and private firms is dependent on the deviation from target trade credit. Conditional on being above the target trade credit level, public firms have a speed of adjustment of 39%, which is higher than the speed of 26% for private firms. In contrast, conditional on being below the target trade credit level, private firms have a higher speed of adjustment than public firms (20% vs. 11%). The former finding suggests that, due to having lower adjustment costs, public firms are able to adjust quickly toward their target trade credit when they are above the target, thus allowing them to mitigate the costs of bankruptcy and the loss of reputation in the

[15] In unreported results, we also follow Gao, Harford, and Li's (2013) approach and assume that public and private firms have the same target trade credit. That is we estimate the target level using a pooled sample of public and private firms. Using this approach, we obtain qualitatively similar results regarding the adjustment speeds.

event of a default. On the other hand, private firms adjust their trade credit more quickly when they are below the target, possibly because of the importance of trade credit as a major source of financing for these firms and hence the need for them to operate closer to the optimal level.

[Insert Table 9 here]

5.4. The use of trade credit during a crisis

We next examine the differential impacts of credit conditions on public and private firms' trade credit polices (*Hypothesis 3*). As mentioned, we use the recent financial crisis of 2007–2009 as a proxy for a credit crunch. Column (1) of Table 10 reports the regression results for model (4). The results show that the crisis has a significantly negative effect on the average level of trade credit for both public and private firms. Additionally, we note that, as in our baseline regression model, the coefficient on the *Public* dummy variable (-0.046) has the expected negative sign, and is statistically and economically significant.

In Column (2), we examine the crisis effects on private and public firms separately. The effect on public firms is captured by the interaction term between *Public* and *Crisis*, plus the stand-alone *Crisis* dummy, which yield a combined effect of 0.2%, or 2% relative to the mean trade credit (9%) of public firms. This effect is statistically significant, but is quite small in magnitude. One possible explanation for this finding is that it captures the combined effects on both constrained and unconstrained public firms. For example, Garcia-Appendini and Montriol-Garriga (2013) find evidence of an increase in the amount of trade credit taken by public firms, but mainly for the constrained group. Another potential explanation is that while Garcia-Appendini and Montriol-Garriga (2013) use quarterly data and focus on the first stage of the crisis before the collapse of Lehman Brothers, from July 2007 to June 2008, we analyze annual data and cover a longer crisis period. This suggests our results could capture the combined

effects of the first-stage of the crisis and the post-Lehman Brothers phase of the crisis associated with a full-blown financial crisis and economic recession.

The impact of the crisis on private firms is captured by the coefficient on *Crisis* (-0.011), which is significantly negative. The coefficient on this dummy variable suggests that during the crisis private firms experienced a one-percent decrease in their trade credit ratio, or a six-percent decrease relative to the mean trade credit of these firms. This finding is not consistent with our conjecture that private firms should use more trade credit during the crisis. However this suggests that although private firms may have demanded more trade credit during the crisis, their demand were not matched by the suppliers' willingness to lend. The finding that supply-side factors may have played an important role in reducing the trade credit of private firms during the crisis is in line with previous research in the literature. Love, Preve, and Sarria-Allende (2007) highlight the importance of the supply of credit during crises. They argue that the increase in trade credit provision in the crisis year, followed by its decline in the post-crisis years, is consistent with a supply-side explanation, as firms vulnerable to the crisis extended less trade credit due to the shortage of funds. Our evidence of a decline in the trade credit of private firms during the crisis is a new and important empirical result, and adds to prior US evidence of the crisis effect on the trade credit of public firms.

In Columns (3)–(4), we further investigate the differential effects of the crisis on the trade credit levels of public and private firms conditional on two measures of financial vulnerability, namely short-term debt and cash flow. Love, Preve, and Sarria-Allende (2007) argue that pre-crisis, firms with a high level of short-term debt (cash flow) are more (less) vulnerable to credit shocks. We consider 2006 as the pre-crisis year and use the 2006 values of those variables as proxies for financial vulnerability. The results in Column (3) show that

conditional on the pre-crisis level of short-term debt, the impact of the crisis on public firms, measured by the sum of the coefficients on *Crisis* (-0.010), *Crisis*×*Public* (0.012), and *Short-term debt*$_{pre\text{-}crisis}$×*Crisis*×*Public* (0.068), is significantly positive and, increasing in *Short-term debt*$_{pre\text{-}crisis}$. This is consistent with Love, Preve, and Sarria-Allende (2007), and suggests that public firms with more short-term debt faced an increased refinancing risk during the crisis, forcing them to rely more on trade credit as a substitute. However, for private firms, the impact of the crisis conditional on the pre-crisis level of short-term debt, measured by the sum of the coefficients on *Crisis* (-0.010) and *Short-term debt*$_{pre\text{-}crisis}$×*Crisis* (-0.085), is significantly negative. This finding is in line with the earlier results in Column (2), and suggests that the more vulnerable private firms were pre-crisis, the more difficult it was for them to obtain trade credit.

Column (4) reports the crisis impact conditional on the pre-crisis level of cash flow. The results in Column (4) show that the impact on public firms, equal to the sum of the coefficients on *Crisis* (-0.007), *Crisis*×*Public* (0.009), and *Cash flow*$_{pre\text{-}crisis}$×*Crisis*×*Public* (0.033), is significantly positive, and increases with the pre-crisis level of cash flow. This finding does not support the argument that internally generated cash flow could be used as a substitute for supplier financing during the crisis. If anything, public firms with more cash flow pre-crisis relied even more on trade credit. However, firms with ample cash flow may signal better creditworthiness and thus have better access to supplier financing. Empirically, our result is not in line with Love, Preve, and Sarria-Allende (2007), who report an insignificant effect of pre-crisis cash flow on the level of trade credit during the Asian and Mexican currency crises. Finally, the impact of the crisis on the trade credit ratio of private firms, equal to the sum of the coefficients on *Crisis* (-0.007) and *Cash flow*$_{pre\text{-}crisis}$×*Crisis* (-0.034), is significantly negative, and increases with the pre-crisis level of cash flow. This suggests that private firms with more

cash flow pre-crisis needed less trade credit during the crisis, which is consistent with the substitution effect.

[Insert Table 10 here]

Overall, we document differential effects of the financial crisis on the trade credit ratios of public and private firms. While public firms used slightly more trade credit during the crisis period, private firms were granted significantly less trade credit. The former finding is consistent with our conjecture that firms substitute short-term borrowings for trade credit during a crisis. However, the latter finding is more in line with a supply-side story whereby firms with a high default risk are likely to be refused trade credit, especially during bad times when the supply of such credit also dries up.

6. Conclusion

Trade credit is one of the most important sources of short-term financing. In this paper, we compare the use of trade credit by public and private firms. We hypothesize that private firms have a higher level of trade credit because they have higher degrees of asymmetric information, financial constraints, and a higher default risk. Using data from the S&P Capital IQ database for the period 1995–2012, we find strong evidence to support this hypothesis. The level of trade credit in private firms is approximately one-quarter higher than that in public firms. This finding is robust to controlling for sample selection and endogeneity concerns.

We also find that the impact of being private on trade credit is most pronounced in young, high-growth, low-tangibility, and large firms. Private firms with young age, high growth opportunities, and low tangibility face higher degrees of asymmetric information and credit constraints, which could explain why these firms rely most heavily on trade credit. Our results

thus show that the difference in public and private firms' use of trade credit varies with the degrees of asymmetric information and constraints facing these firms.

We further examine optimal trade credit policies for both public and private firms. Our findings show that both firm types adjust quite actively toward their target trade credit. However, public firms are able to move faster to their target with an adjustment speed of 29%, compared to a speed of 23% for private firms. This is consistent with the argument that public firms have lower adjustment costs due to their greater bargaining power with suppliers as well as greater access to other forms of credit. While our findings show how reliant private firms are on trade credit, they also indicate how relatively difficult it is for these firms to adjust their trade credit and maintain its level close to, or at, the optimum.

Finally, we document differential effects of credit conditions on the use of trade credit by public and private firms. During the recent financial crisis, public firms were able to use slightly more trade credit as a substitute for other sources of financing, while private firms were granted significantly less trade credit and were forced to rely on internally generated funds. The latter finding suggests that a supply-side story may be at work as the demand for more trade credit by private firms may not be accommodated by suppliers during a credit crunch when both bank and trade credit become less accessible. Overall, while our results show the importance of trade credit to private firms, they also highlight the limitations of this form of credit in absorbing credit shocks. Our study, thus, provides implications for policies aimed at enhancing the flow of lending to private firms during times of extreme financial stress.

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Figure 1: Evolution of Trade Credit and the Listing Status

This figure illustrates the evolution of trade credit around the IPO year for a transition sample of 1,282 IPO firms that went public during the sample period. Trade credit is defined as the ratio of accounts payable to total assets.

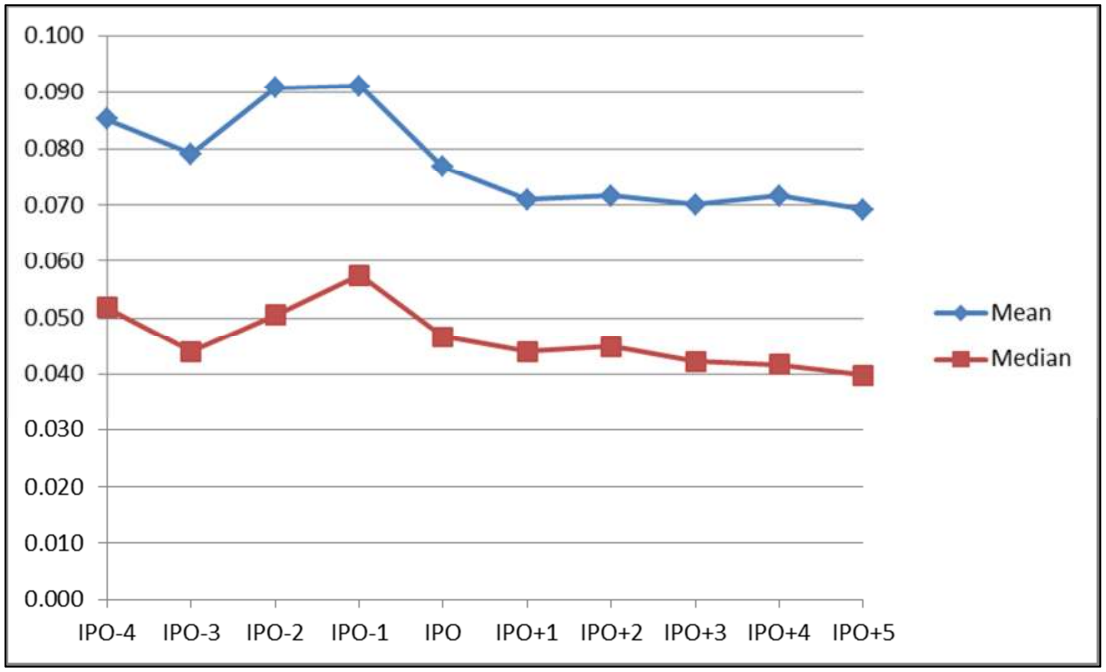

Table 1: Summary Statistics and Univariate Analysis

Panel A of this table reports the summary statistics of the variables under consideration. Trade credit is defined as the ratio of accounts payable to total assets. Firm age is the number of years from incorporation. Cash flow is measured by net profits plus depreciation, scaled by total assets. Cash holdings is cash and cash equivalents, scaled by total assets. Current assets is current assets minus cash, scaled by total assets. Negative growth is sales growth times the negative growth dummy variable, which takes the value of 1 if sales growth is negative, and 0 otherwise. Positive growth is sales growth times the positive growth dummy variable, which takes the value of 1 if sales growth is positive, and 0 otherwise. Short-term debt is short-term borrowings plus the current portion of long-term debt, scaled by total assets. Size is the natural logarithm of total sales, measured in 2012 dollar price. The total number of observations in our sample is 103,777. Panel B presents the univariate analysis of the trade credit of the public and private firms in the full sample using the t-test for differences in mean and the Wilcoxon-Mann-Whitney test for differences in median.

Panel A: Summary Statistics

Variable	Mean	Median	Std dev	Min	Max
Trade credit (%)	0.1359	0.0898	0.1324	0.0077	0.5612
Age (years)	39.97	31.000	29.49	0.0000	135.00
Cash flow (%)	0.0861	0.0540	0.2001	-0.5000	1.5003
Cash holdings (%)	0.1344	0.7742	0.1553	0.0000	1.0000
Current assets (%)	0.5071	0.5098	0.2561	0.0000	0.9819
Negative growth	-0.0523	0.0000	0.1262	-1.0000	0.0000
Positive growth	0.2068	0.0650	0.7545	0.0000	15.7500
Short-term debt (%)	0.0301	0.0000	0.0843	0.0000	1.0000
Size (Ln)	2.7121	2.1207	2.0889	-3.6498	8.2483

Panel B: Univariate Analysis

Trade credit	Public	Private	Mean test (p-value)	Median test (p-value)
Mean	0.0896	0.1582	0.000	
Median	0.0663	0.1108		0.000
Number of obs	70,011	33,766		

Table 2: The Use of Trade Credit by Public and Private Firms across Industries

This table shows the level of trade credit according to the 12-Fama French industry classification. Trade credit is defined as the ratio of accounts payable to total assets. Panel A provides the summary statistics (i.e., the mean, median, standard deviation (std. dev.), minimum, and maximum), of the trade credit ratio across industries. Panel B shows how the trade credit of public and private firms varies across industries. It reports the p-values of the t-test of differences in mean (i.e., mean test) and the p-value of the Wilcoxon-Mann-Whitney test of differences in median (median test).

Panel A: Full Sample

Industry	Industry description	Mean	Median	Std. dev.	Min	Max	N. obs.
1	Consumer non-durables	0.1064	0.0787	0.0941	0.0109	0.4732	4,794
2	Consumer durables	0.1257	0.0998	0.0979	0.0147	0.5621	1,866
3	Manufacturing	0.1080	0.0863	0.0797	0.0171	0.3829	11,762
4	Energy	0.0819	0.0548	0.0936	0.0019	0.9086	2,536
5	Chemicals	0.1049	0.0880	0.0779	0.0159	0.6777	1,891
6	Business equipment	0.1018	0.0637	0.1140	0.0069	0.7944	9,787
7	Telecommunications	0.0690	0.0337	0.1083	0.0029	0.7876	1,610
9	Retail and wholesale	0.1882	0.1474	0.1473	0.0168	0.5751	26,823
10	Health	0.0654	0.0409	0.0758	0.0077	0.5261	9,098
12	Others	0.1467	0.0938	0.1451	0.0054	0.5507	33,610

Panel B: Public Firms versus Private Firms

Industry	Industry description	Firm type	Mean	Median	Std. dev.	N. obs.	Mean test	Median test
1	Consumer non-durables	Private	0.1295	0.0915	0.1149	2,383		
		Public	0.0836	0.0715	0.0591	2,411	0.000	0.000
2	Consumer durables	Private	0.1480	0.1182	0.1217	710		
		Public	0.1120	0.0940	0.0767	1,156	0.000	0.000
3	Manufacturing	Private	0.1226	0.0969	0.0933	6,460		
		Public	0.0904	0.0788	0.0539	5,302	0.000	0.000
4	Energy	Private	0.1517	0.0887	0.1717	340		
		Public	0.0711	0.0524	0.0685	2,196	0.000	0.000
5	Chemicals	Private	0.1330	0.1035	0.1132	523		
		Public	0.0942	0.0836	0.0554	1,368	0.000	0.000
6	Business equipment	Private	0.1547	0.0982	0.1604	2,846		
		Public	0.0800	0.0564	0.0784	6,941	0.000	0.000
7	Telecommunications	Private	0.0776	0.0305	0.1282	708		
		Public	0.0622	0.0365	0.0892	902	0.005	0.048
9	Retail and wholesale	Private	0.1949	0.1535	0.1516	22,440		
		Public	0.1541	0.1238	0.1178	4,383	0.000	0.000
10	Health	Private	0.0695	0.0394	0.0861	5,058		
		Public	0.0603	0.0426	0.0602	4,040	0.000	0.370
12	Others	Private	0.1587	0.1076	0.1495	28,543		
		Public	0.0793	0.0479	0.0922	5,067	0.000	0.000

Table 3: Difference in Trade Credit between Public and Private Firms

This table reports the regression results for model (1), which captures the difference in the trade credit of public and private firms. Columns (1)–(4) provide the results for the full sample. Column (5) provides the results for a matched sample, in which we match, with replacement, each public firm to any private firm in the same industry and year, and of closest size (allowing for a deviation of 30%). The dependent variable is trade credit, defined as the ratio of accounts payable to total assets. *Public* is a dummy variable that takes the value of 1 for public firms, and 0 otherwise. Firm age is the number of years from incorporation. Cash flow is net profits plus depreciation, scaled by total assets. Cash holdings is cash and cash equivalents, scaled by total assets. Current assets is current assets minus cash, scaled by total assets. Negative (positive) growth is sales growth times the negative (positive) growth dummy, which takes the value of 1 if sales growth is negative (positive), and 0 otherwise. Short-term debt is short-term borrowings plus the current portion of long-term debt, scaled by total assets. Size is the natural logarithm of total sales, measured in 2012 dollar price. T-statistics are reported in parentheses. Standard errors are heteroskedasticity robust. ***, **, and * denote statistical significance at the 1, 5, and 10 percent levels, respectively.

	(1)	(2)	(3)	(4)	(5)
Public	-0.0686***	-0.0435***	-0.0554***	-0.0362***	-0.0659***
	(-96.26)	(-45.56)	(-60.04)	(-31.66)	(-53.97)
ln(1+age)			-0.0221***	-0.0198***	-0.0046
			(-8.93)	(-8.07)	(-1.16)
ln(1+age)^2			0.0004	-0.0002	-0.00197***
			(1.13)	(-0.50)	(-3.38)
Cash flow			-0.0471***	-0.0474***	-0.0644***
			(-21.86)	(-22.15)	(-27.59)
Cash holdings			0.0770***	0.0761***	0.1043***
			(29.03)	(27.75)	(32.06)
Current assets			0.2228***	0.1961***	0.1912***
			(146.43)	(103.32)	(90.72)
Negative growth			0.0099***	0.0224***	0.0133***
			(3.18)	(7.08)	(3.84)
Positive growth			0.0048***	0.0039***	0.0059***
			(8.98)	(8.04)	(4.59)
Short-term debt			-0.0646***	-0.0707***	-0.0619***
			(-14.65)	(-16.14)	(-12.20)
Size			0.0090***	0.0079***	0.0094***
			(40.79)	(36.44)	(25.19)
Intercept			0.0825***	0.0817***	0.0405***
			(19.14)	(14.12)	(3.42)
Industry effects	No	Yes	No	Yes	Yes
Year effects	No	Yes	No	Yes	Yes
Number of obs.	103,777	103,777	103,777	103,777	84,243
Adjusted R^2	0.059	0.176	0.236	0.283	0.268

Table 4: Regression Results for the Transition Sample

This table presents the regression results for the transition sample of 1,282 IPO firms. The dependent variable is trade credit, defined as the ratio of accounts payable to total assets. D_{Post_IPO} is a dummy variable that takes the value of 1 from the IPO year onwards, and 0 in the years pre-IPO. D_{IPO+i} with i=1..3 is a dummy variable that takes the value of 1 in the IPO+i year, and 0 otherwise. The other independent variables are defined in Table 3. T-statistics are reported in parentheses. Standard errors are heteroskedasticity robust. ***, **, and * denote statistical significance at the 1, 5, and 10 percent levels, respectively.

	(1)	(2)	(3)	(4)	(5)	(6)
D_{Post_IPO}	-0.0223***	-0.0148***	-0.0113***			
	(-9.48)	(-6.55)	(-4.21)			
D_{IPO}				-0.0087**	-0.0104***	-0.0111***
				(-2.35)	(-3.06)	(-3.35)
D_{IPO+1}				-0.0169***	-0.0150***	-0.0148***
				(-4.85)	(-4.38)	(-3.57)
D_{IPO+2}				-0.0176***	-0.0133***	-0.0129**
				(-4.93)	(-3.85)	(-2.45)
D_{IPO+3}				-0.0207***	-0.0144***	-0.0174***
				(-5.54)	(-4.02)	(-2.66)
ln(1+age)	-0.0199***	-0.0055	-0.0099	-0.0150***	-0.0039	0.0049
	(-5.20)	(-1.50)	(-1.18)	(-3.03)	(-0.86)	(0.51)
$\ln(1+age)^2$	0.0041***	0.0014**	0.0019	0.0032***	0.0009	-0.0009
	(6.14)	(2.23)	(0.53)	(3.45)	(1.02)	(-0.14)
Cash flow	-0.0064	-0.0265***	-0.0069	0.0007	-0.0277***	-0.0023
	(-1.08)	(-4.48)	(-1.22)	(0.07)	(-2.97)	(-0.24)
Cash holdings	0.01685***	0.04081***	0.0215**	0.0103	0.0309***	-0.0185
	(3.56)	(8.23)	(2.42)	(1.63)	(4.57)	(-1.64)
Current assets	0.1394***	0.1357***	0.0844***	0.1405***	0.1310***	0.0493***
	(30.34)	(26.86)	(7.09)	(20.62)	(17.54)	(3.74)
Negative growth	0.0036	0.0049	0.0166**	-0.0001	0.0015	0.0244***
	(0.45)	(0.61)	(2.40)	(-0.01)	(0.12)	(2.64)
Positive growth	-0.0002	-0.0002	8.34×10^{-05}	0.0001	-1.18×10^{-05}	0.0003
	(-0.33)	(-0.35)	(0.20)	(0.17)	(-0.02)	(0.70)
Short-term debt	0.0896***	0.0086	0.0525***	0.0826***	0.0067	0.0620***
	(6.73)	(0.66)	(3.55)	(3.76)	(0.30)	(2.72)
Size	0.0061***	0.0028***	-9.84×10^{-05}	0.0076***	0.0042***	-0.0024
	(9.35)	(4.46)	(-0.06)	(7.59)	(4.10)	(-1.00)
Intercept	0.0334***	0.0319*	0.0894***	0.0182**	0.0209	0.0786*
	(5.59)	(1.84)	(5.81)	(2.36)	(1.17)	(1.67)
Industry effects	No	Yes	Yes	No	Yes	Yes
Year effects	No	Yes	Yes	No	Yes	Yes
Firm effects	No	No	Yes	No	No	Yes
Number of obs.	9,976	9,976	9,976	4,444	4,444	4,444
Adjusted R^2	0.170	0.302	0.086	0.181	0.332	0.127

Table 5: Treatment Regression Results

This table reports the results of the treatment regression models (6) and (7), in which the *Public* dummy variable, which takes the value of 1 for public firms, and 0 otherwise, is assumed to be endogenous. Columns (1)–(2) report the first- second-stage regression results, respectively. The first-stage regression is a probit regression with the *Public* dummy being the dependent variable. We instrument for this variable using the industry underwriter concentration variable, defined as the number of IPOs underwritten by the Top 5 underwriters divided by the number of IPOs in that industry. Here, the Top 5 underwriters are determined by the number of IPOs that they have underwritten over the last five years. The other independent variables are defined in Table 3. T-statistics are reported in parentheses. Standard errors are heteroskedasticity robust. ***, **, and * denote statistical significance at the 1, 5, and 10 percent levels, respectively.

	(1)	(2)
Public		-0.0702***
		(-10.94)
ln(1+age)	-0.0450	-0.0228***
	(-1.14)	(-8.54)
$\ln(1+age)^2$	-0.0372***	0.0004
	(-6.25)	(1.10)
Cash flow	-2.5536***	-0.0506***
	(-47.80)	(-15.25)
Cash holdings	0.2808***	0.0766***
	(7.18)	(27.18)
Current assets	-1.4988***	0.2167***
	(-63.96)	(79.56)
Negative growth	-1.1543***	0.0065*
	(-23.09)	(1.68)
Positive growth	0.5485***	0.0055***
	(28.02)	(7.96)
Short-term debt	1.8264***	-0.0603***
	(28.88)	(-10.78)
Size	0.4584***	0.0109***
	(119.46)	(12.88)
Underwriter concentration	-0.3796***	
	(-22.29)	
Intercept	-0.4114***	0.0873***
	(-6.04)	(16.83)
Endogeneity test		0.0817**
		(2.49)
Number of observations		92,495

Table 6: Propensity Score Matching

This table reports the propensity score matching results. We match each public firm to a private firm using propensity score matching to the nearest neighborhood, without replacement. In Specification (1), the matching is based on size, industry, and year. In (2), the matching is based on all the control variables, namely ln(1+age), $\ln(1+age)^2$, cash holdings, cash flow, current assets, negative growth, positive growth, short-term debt, size, industry, and year effects. Panel A presents the pairwise differences in the mean trade credit in the matched samples. Bootstrapped standard errors (Std. error) based on 50 replications are reported in the square brackets. Panel B reports the baseline regression results for the matched samples. T-statistics are reported in parentheses. Standard errors are heteroskedasticity robust. ***, **, and * denote statistical significance at the 1, 5, and 10 percent levels, respectively.

Panel A: Pairwise differences in the mean trade credit

	Specification (1)	**Specification (2)**
Difference	-0.0793***	-0.0607***
Std. error	[0.0008]	[0.0011]

Panel B: Regression results for the propensity-score matched sample

	Specification (1)	**Specification (2)**
Public	-0.0369***	-0.0359***
	(-28.94)	(-30.04)
ln(1+age)	-0.0159***	-0.0162***
	(-6.36)	(-6.78)
$\ln(1+age)^2$	-1.40×10^{-05}	0.0003
	(-0.04)	(0.74)
Cash flow	-0.0429***	-0.0703***
	(-18.92)	(-21.24)
Cash holdings	0.0704***	0.0736***
	(22.08)	(26.38)
Current assets	0.1896***	0.2049***
	(86.93)	(97.57)
Negative growth	0.0219***	0.0381***
	(5.85)	(11.96)
Positive growth	0.0026***	0.0024***
	(5.52)	(5.38)
Short-term debt	-0.0520***	-0.0826***
	(-10.42)	(-19.04)
Size	0.0054***	0.0054***
	(22.25)	(22.33)
Intercept	0.0748***	0.0729***
	(11.34)	(11.41)
Industry effects	Yes	Yes
Year effects	Yes	Yes
Number of observations	67,532	67,532
Adjusted R^2	0.322	0.325

Table 7: Additional Robustness Checks

This table presents the regression results from four robustness checks. Columns (1)-(2) report the results after including additional control variables. Specifically, Column (1) includes Fcost, which is interest expense divided by the sum of total debt minus accounts payable, while Column (2) includes the annual rate of GDP growth. Column (3) presents the regression results for an extended sample including utilities firms. Column (4) reports the regression results in which an alternative definition of the dependent variable (accounts payable over cost of goods sold) is used. The other independent variables are defined in Table 3. T-statistics are reported in parentheses. Standard errors are heteroskedasticity robust. ***, **, and * denote statistical significance at the 1, 5, and 10 percent levels, respectively.

	(1)	(2)	(3)	(4)
Public	-0.0383***	-0.0362***	-0.0346***	-0.0219***
	(-29.61)	(-31.66)	(-31.68)	(-2.04)
ln(1+age)	-0.0197***	-0.0198***	-0.0202***	-0.0646***
	(-8.03)	(-8.07)	(-8.44)	(-5.86)
$\ln(1+age)^2$	-0.0002	-0.0002	3.21×10^{-05}	0.0083***
	(-0.48)	(-0.50)	(0.09)	(5.26)
Cash flow	-0.0469***	-0.0474***	-0.0461***	-0.1139***
	(-21.92)	(-22.15)	(-21.75)	(-6.94)
Cash holdings	0.0769***	0.0761***	0.0758***	-0.0902***
	(27.83)	(27.75)	(27.96)	(-5.13)
Current assets	0.1967***	0.1961***	0.1966***	-0.1558***
	(103.20)	(103.32)	(104.80)	(-14.89)
Negative growth	0.0224***	0.0224***	0.0176***	-0.3150***
	(7.08)	(7.08)	(8.08)	(-11.19)
Positive growth	0.0039***	0.0039***	0.0042***	0.0249***
	(7.74)	(8.04)	(8.62)	(8.49)
Short-term debt	-0.0724***	-0.0707***	-0.0687***	0.0178
	(-16.46)	(-16.14)	(-15.79)	(0.76)
Size	0.0081***	0.0079***	0.0078***	-0.0209***
	(36.68)	(36.44)	(37.57)	(-19.30)
Fcost	0.0564***			
	(3.26)			
ΔGDP		0.0005*		
		(1.72)		
Intercept	0.0805***	0.0739***	0.0759***	0.3617***
	(13.88)	(13.23)	(13.63)	(13.52)
Industry effects	Yes	Yes	Yes	Yes
Year effects	Yes	Yes	Yes	Yes
Number of obs.	103,777	103,777	107,338	33,922
Adjusted R^2	0.283	0.283	0.295	0.238

Table 8: The Use of Trade Credit by Public and Private Firms Conditional on Firm Characteristics

This table presents the effect of four firm-specific characteristics, namely firm age, growth opportunities, tangibility, and firm size, on the difference in the level of trade credit used by public and private firms. The dependent variable is trade credit, defined as the ratio of accounts payable to total assets. *Mature* is a dummy variable that takes the value of 1 for observations with above the median age, and 0 otherwise. *High growth* is a dummy variable that takes the value of 1 for observations with above the median growth opportunities, and 0 otherwise. *High tangibility* is a dummy variable that takes the value of 1 for observations with above the median tangibility, and 0 otherwise. *Large* is a dummy variable that takes the value of 1 for observations with above the median size, and 0 otherwise. *Public* is a dummy variable that takes the value of 1 for public firms, and 0 otherwise. The other independent variables are defined in Table 3. T-statistics are reported in parentheses. Standard errors are heteroskedasticity robust. ***, **, and * denote statistical significance at the 1, 5, and 10 percent levels, respectively.

Age (1)		**Growth opportunities** (2)		**Tangibility** (3)		**Size** (4)	
Mature	-0.0199***	High growth	0.0178***	High tangibility	-0.0164***	Large	0.0127***
	(-15.35)		(17.43)		(-15.15)		(10.90)
Public	-0.0492***	Public	-0.0274***	Public	-0.0532***	Public	-0.0133***
	(-35.91)		(-21.41)		(-35.21)		(-8.24)
Mature×Public	0.0272***	High growth×Public	-0.0178***	High tangibility×Public	0.0311***	Large×Public	-0.0349***
	(20.15)		(-13.90)		(21.98)		(-19.52)
ln(1+age)	-0.0195***	ln(1+age)	-0.0191***	ln(1+age)	-0.0188***	ln(1+age)	-0.0189***
	(-8.10)		(-7.84)		(-7.73)		(-7.73)
$\ln(1+age)^2$	0.0006*	$\ln(1+age)^2$	-0.0002	$\ln(1+age)^2$	-0.0003	$\ln(1+age)^2$	-0.0003
	(1.72)		(-0.49)		(-0.87)		(-0.74)
Cash flow	-0.0507***	Cash flow	-0.0498***	Cash flow	-0.0491***	Cash flow	-0.0443***
	(-23.44)		(-23.03)		(-22.81)		(-20.73)
Cash holdings	0.0763***	Cash holdings	0.0765***	Cash holdings	0.0656***	Cash holdings	0.0743***
	(27.85)		(27.94)		(23.46)		(27.22)
Current assets	0.1940***	Current assets	0.1955***	Current assets	0.1829***	Current assets	0.1959***
	(102.01)		(103.17)		(91.10)		(103.30)
Negative growth	0.0239***	Negative growth	0.0054	Negative growth	0.0239***	Negative growth	0.0231***
	(7.54)		(1.62)		(7.61)		(7.34)
Positive growth	0.0044***	Positive growth	0.0028***	Positive growth	0.0041***	Positive growth	0.0031***
	(8.86)		(5.90)		(8.39)		(6.41)
Short-term debt	-0.0689***	Short-term debt	-0.0716***	Short-term debt	-0.0709***	Short-term debt	-0.0748***
	(-15.73)		(-16.36)		(-16.14)		(-17.11)
Size	0.0072***	Size	0.0079***	Size	0.0075***	Size	0.0083***
	(32.48)		(35.97)		(34.37)		(29.44)
Intercept	0.0852***	Intercept	0.0714***	Intercept	0.0979***	Intercept	0.0757***
	(14.67)		(12.32)		(16.63)		(13.08)
Industry effects	Yes	Industry effects	Yes	Industry effects	Yes	Industry effects	Yes
Year effects	Yes	Year effects	Yes	Year effects	Yes	Year effects	Yes
Number of obs.	103,777	Number of obs.	103,777	Number of obs.	103,777	Number of obs.	103,777
Adjusted R^2	0.286	Adjusted R^2	0.285	Adjusted R^2	0.286	Adjusted R^2	0.285

Table 9: Speed of Adjustment to Target Trade Credit

This table presents the regression results for the partial adjustment model of trade credit (2). It reports the estimated speed of adjustment, showing how fast public and private firms adjust toward their respective target level of trade credit. Panel A presents the results for the full sample of public and private firms. Panel B provides the results for the subsample of public and private firms with above-target trade credit. Panel C reports the results for the subsample of public and private firms with below-target trade credit. The dependent variable, ΔTC_{it}, is the change in trade credit. The independent variable, $TC_{it}{}^{*} - TC_{it-1}$, is the deviation from target trade credit, where $TC_{it}{}^{*}$ is the estimated target trade credit (see also Table A-2 in the Appendix for more details). P-values of the Chow test of differences in the adjustment speed estimates are reported in square brackets. T-statistics are reported in parentheses. Standard errors are heteroskedasticity robust. ***, **, and * denote statistical significance at the 1, 5, and 10 percent levels, respectively.

Panel A: Full Sample

	Public Firms ΔTC_{it}	Private Firms ΔTC_{it}	F-stat of Chow test [p-value]
$TC_{it}{}^{*} - TC_{it-1}$	0.2919***	0.2316***	38.32
	(97.85)	(101.74)	[0.000]
Intercept	-0.0029***	-0.0018***	
	(-11.95)	(-6.11)	
Number of observations	33,766	70,011	
Adjusted R^2	0.221	0.129	

Panel B: Firms with Above-target Trade Credit

	Public Firms ΔTC_{it}	Private Firms ΔTC_{it}	F-stat of Chow test [p-value]
$TC_{it}{}^{*} - TC_{it-1}$	0.3956***	0.2638***	66.53
	(68.88)	(53.13)	[0.000]
Intercept	0.0088***	0.0032***	
	(14.16)	(4.10)	
Number of observations	13,929	29,459	
Adjusted R^2	0.254	0.087	

Panel C: Firms with Below-target Trade Credit

	Public Firms ΔTC_{it}	Private Firms ΔTC_{it}	F-stat of Chow test [p-value]
$TC_{it}{}^{*} - TC_{it-1}$	0.1088***	0.2046***	68.65
	(17.36)	(36.24)	[0.000]
Intercept	0.0011***	-0.0005	
	(3.54)	(-0.82)	
Number of observations	19,837	40,552	
Adjusted R^2	0.015	0.031	

Table 10: The Effects of the Financial Crisis on the Use of Trade Credit by Public and Private Firms

This table presents the regression results for model (4), which captures the effects of the recent financial crisis on the trade credit ratios of public and private firms. The model is estimated for the period 2004–2009. The crisis period is defined as from 2007 to 2009. The dependent variable is trade credit, defined as the ratio of accounts payable to total assets. *Crisis* is a dummy variable that takes the value of 1 for the years 2007–2009, and 0 otherwise. *Public* is a dummy variable that takes the value of 1 for public firms, and 0 otherwise. *Short-term debt*$_{pre\text{-}crisis}$ is the pre-crisis (2006) level of short-term debt. *Cash flow*$_{pre\text{-}crisis}$ is the pre-crisis (2006) level of cash flow. The model includes all the control variables, listed and defined as in Table 3. T-statistics are reported in parentheses. Standard errors are heteroskedasticity robust. ***, **, and * denote statistical significance at the 1, 5, and 10 percent levels, respectively.

	(1)	(2)	(3)	(4)
Crisis	-0.0082***	-0.0106***	-0.0099***	-0.0072***
	(-4.68)	(-5.71)	(-5.35)	(-3.81)
Public	-0.0402***	-0.0464***	-0.0524***	-0.0461***
	(-24.53)	(-24.44)	(-26.36)	(-24.29)
Crisis*Public		0.0127***	0.0122***	0.0088***
		(6.56)	(5.91)	(4.41)
Short-term deb$_{pre\text{-}crisis}$×Crisis			-0.0849***	
			(-6.13)	
Short-term debt$_{pre\text{-}crisis}$×Public			0.1288***	
			(7.19)	
Short-term debt$_{pre\text{-}crisis}$×Crisis×Public			0.0681**	
			(2.50)	
Cash flow$_{pre\text{-}crisis}$×Crisis				-0.0341***
				(-8.92)
Cash flow$_{pre\text{-}crisis}$×Public				0.0002
				(0.39)
Cash flow$_{pre\text{-}crisis}$×Crisis×Public				0.0331***
				(8.26)
Controls	Yes	Yes	Yes	Yes
Industry effects	Yes	Yes	Yes	Yes
Year effects	Yes	Yes	Yes	Yes
Number of observations	56,283	56,283	56,283	56,283
Adjusted R^2	0.273	0.273	0.274	0.274

Appendix

Table A.1. Comparison between data on trade credit provided by S&P Capital IQ and Compustat

This table provides a comparison of the summary statistics (i.e., the mean, median, and standard deviation) of the trade credit of the public firms in our sample and the public firms in the Compustat database. Trade credit is defined as the ratio of accounts payable to total assets.

Trade Credit	Our Sample	Compustat
Mean	0.0895	0.0894
Median	0.0663	0.0642
Std dev.	0.0836	0.0882
N	33,766	114,845

Table A.2. Estimation of the Target Trade Credit Levels of Public and Private Firms

This table reports the regression results for the estimation of the target level trade credit as given by model (3). Trade credit is defined as the ratio of accounts payable to total assets. The other independent variables are defined in Table 3. P-values of the Chow test for differences in the coefficient estimates are reported in square brackets. T-statistics are reported in parentheses. Standard errors are heteroskedasticity robust. ***, **, and * denote statistical significance at the 1, 5, and 10 percent levels, respectively.

	Public firms	**Private firms**	**F-stat of Chow test [p-value]**
ln(1+age)	-0.01758***	-0.006344	5.93
	(-6.88)	(-1.65)	[0.015]
ln(1+age)^2	0.001911***	-0.003041***	52.59
	(5.11)	(-5.32)	[0.000]
Cash flow	-0.05135***	-0.05404***	0.33
	(-13.11)	(-21.52)	[0.564]
Cash holdings	0.001920	0.1133***	562.55
	(0.63)	(31.69)	[0.000]
Current assets	0.1362***	0.2197***	478.70
	(45.00)	(94.01)	[0.000]
Negative growth	0.01770***	0.01459***	0.29
	(4.49)	(3.44)	[0.592]
Positive growth	0.0007084	0.03145***	193.96
	(1.62)	(14.53)	[0.000]
Short-term debt	0.05960***	-0.1397***	565.98
	(9.04)	(-26.96)	[0.000]
Size	0.001856***	0.01432***	641.06
	(8.18)	(32.74)	[0.000]
Intercept	0.04133***	0.02501	
	(6.50)	(1.56)	
Industry effects	Yes	Yes	
Year effects	Yes	Yes	
N	103,777	103,777	
Adjusted R^2	0.262	0.260	