# CEB Working Paper

Centre Emile Bernheim
Research Institute in Management Sciences

Solvay Brussels School
Economics & Management

# Trade Credit and Product Market Power during a Financial Crisis

**Adalto Barbaceia Gonçalves, Rafael Schiozer and Hsia Hua Sheng**

This paper investigates whether product market power affects trade credit decisions. We exploit the 2007-08 credit crisis in the U.S. as a source of variation in the importance of product market power for trade credit. We find that a one standard deviation increase in market power is associated to a decrease in payables of approximately four days during the crisis, showing that high market power firms alleviate financial constraints from their suppliers to avoid the loss of monopoly rents. Our inferences are robust to structural and non-structural measures of market power, both at the firm and at the industry levels, and the inclusion of controls to address potential confounding effects deriving from other firm features, including financial constraints, industry specific shocks and macroeconomic effects.

Keywords: trade credit, financial crisis, market power, monopoly rents, liquidity provision.

JEL Classifications: G01; G30; G32; D43.

CEB Working Paper N° 18/004
January 2018

*Université Libre de Bruxelles - Solvay Brussels School of Economics and Management*
*Centre Emile Bernheim*
*ULB CP114/03 50, avenue F.D. Roosevelt 1050 Brussels BELGIUM*
*e-mail: ceb@admin.ulb.ac.be Tel.: +32 (0)2/650.48.64 Fax: +32 (0)2/650.41.88*

ULB

# Trade Credit and Product Market Power during a Financial Crisis

*ADALTO BARBACEIA GONÇALVES*[1]

*RAFAEL SCHIOZER*[2]

*HSIA HUA SHENG*[3]

**Abstract:** *This paper investigates whether product market power affects trade credit decisions. We exploit the 2007-08 credit crisis in the U.S. as a source of variation in the importance of product market power for trade credit. We find that a one standard deviation increase in market power is associated to a decrease in payables of approximately four days during the crisis, showing that high market power firms alleviate financial constraints from their suppliers to avoid the loss of monopoly rents. Our inferences are robust to structural and non-structural measures of market power, both at the firm and at the industry levels, and the inclusion of controls to address potential confounding effects deriving from other firm features, including financial constraints, industry specific shocks and macroeconomic effects.*

**Keywords:** trade credit, financial crisis, market power, monopoly rents, liquidity provision

**JEL classifications**: G01; G30; G32; D43

[1] *Insper Institute of Education, São Paulo – Brazil. Email: adaltobg@insper.edu.br*
[2] ***Corresponding author**: Fundação Getulio Vargas / EAESP – Email: rafael.schiozer@fgv.br; Av. Nove de Julho 2029, Bela Vista, 01313-902 - São Paulo – SP – Brazil; Phone: (+55)11 37997899*
[3] *Fundação Getulio Vargas / EAESP - hsia.sheng@fgv.br*

# Trade Credit and Market Power during a Financial Crisis

**Abstract:** *This paper investigates whether product market power affects trade credit decisions. We exploit the 2007-08 credit crisis in the U.S. as a source of variation in the importance of product market power for trade credit. We find that a one standard deviation increase in market power is associated to a decrease in payables of approximately four days during the crisis, showing that high market power firms alleviate financial constraints from their suppliers to avoid the loss of monopoly rents. Our inferences are robust to the use of structural and non-structural measures of market power, both at the firm and at the industry levels, and to the inclusion of controls to address potential confounding effects deriving from other firm features, including financial constraints, industry specific shocks and macroeconomic effects.*

**Keywords:** trade credit, financial crisis, market power, monopoly rents, liquidity provision

**JEL classifications**: G01; G30; G32; D43

# 1. INTRODUCTION

The dependency relation between seller and buyer affects the extension and uptake of trade credit by firms. There are opposing theoretical predictions regarding market power and the supply and demand of trade credit. Originally, Meltzer (1960) stated that greater market power implies more trade credit extension. Petersen and Rajan (1997) have empirically confirmed his predictions. Biais and Gollier (1997), Wilner (2000) and Cuñat (2007) have also developed theoretical models predicting that suppliers with more market power extend more trade credit. However, a set of recent papers [Wilson and Summers (2002), Fabbri and Klapper (2008), Giannetti, Burkart, and Ellingsen (2011)] predicts exactly the opposite. According to these studies, small firms with low bargaining power sell with low margins to large customers, supply more trade credit, and even tolerate payment delays. Low market power firms also offer relatively more early payment discounts to get their products' quality inspected and at the same time receive more liquidity from their higher market power customers [Giannetti et al. (2011), Klapper et al. (2012) and Dass, Kale, and Nanda (2015)]. Additionally, the model by Barbosa, Moreira, and Novaes (2017) implies that interest rates embedded in trade credit discounts increase for suppliers that operate with lower margins (low market power).

This paper studies how product market power affects trade credit decisions of US firms using the 2007/08 financial crisis as a source of variation in the importance of product market power for trade credit decisions. We therefore treat market power as a *latent firm feature that becomes more important during a financial crisis*. As argued by Garcia-Appendini and Montoriol-Garriga (2013), this crisis provides a unique laboratory in the investigation of the behavior of non-financial firms, because it originated mostly in the housing and financial sector. Therefore, the reduction in bank loans and capital markets funding to firms was unexpected, and its origins were almost unrelated to nonfinancial firms.

Firms with higher market power enjoy monopolistic rents, and any disruptions along the supply chain could threaten the maintenance of these rents. For example, if suppliers of high market power firms enter into financial distress because of the credit crunch and have difficulty in meeting orders, they could interrupt production along the supply chain and harm the monopoly rents of downstream firms. This idea is consistent with Silva, Ribeiro, and Sheng (2011) and Daripa and Nilsen (2011) who show that delays from suppliers create an externality cost of lost sales. Mateut (2014) confirms that it may be optimal for monopolistic firms to anticipate payments to avoid suppliers from delaying production. Combining both ideas, we claim that the cost of lost sales due to disruptions in production is larger for high than for low market power firms. Because the credit crisis likely changes the probability that a supplier will enter into financial distress, it affects high and low market power firms' trade credit decisions heterogeneously, and therefore high market power firms may be more likely to anticipate payments to suppliers during a crisis.

High market power is also associated with high-quality differentiated products and services. If upstream firms reduce the quality of goods and services supplied because of financial distress, high market power firms may also lose their ability to charge higher prices than their competitors and will lose monopoly rents. Therefore, this is yet another reason why high market power firms may want to alleviate the financial constraints of their suppliers during a financial crisis, and they may decide to do so by reducing the time to pay for goods and services provided (i.e., reduce payable days).

Our results are consistent with the rationale presented above. Our sample uses Compustat data on U.S. firms from 2004 to 2010. Our main measure of market power in the product market is the Lerner index, which is measured by the price-cost margin (PCM) proxied by the firms´ gross margin, following Petersen and Rajan (1997). The Lerner index is widely used as a competition measure by empirical studies, as buyer concentration tends to be negatively correlated with margins [Collins and Preston (1969), Schmalensee (1989), Sutton (2007) and Dass, Kale, and Nanda (2015)].

We find that the larger the firm's market power, the more it reduces payment time to suppliers after the eruption of the crisis, in comparison to the pre-crisis period. A one standard deviation increase in pre-crisis market power is associated to a supplier payment period (payable days) that is 16 days longer. These results confirm Long et al.'s (1993) and Giannetti et al.'s (2011) theories of product quality asymmetry, as high market power firms may require more time from their suppliers to verify the product. During the crisis, however, this effect is reduced by approximately 4 days, showing that high market power firms anticipate payments to their suppliers during the crisis. This anticipation is statistically and economically significant and represents approximately 10% of the median firm payable days in our sample. These findings are compliant with the optimality in anticipating payments to suppliers to avoid delays and disruptions in the supply of goods and services [Daripa and Nilsen (2011)], particularly by firms enjoying monopoly rents. The results are also consistent with the larger early payment discounts offered to high market power firms by firms with low bargaining power [Giannetti et al. (2011), Klapper et al. (2012)].

Before the crisis, a one standard deviation increase in pre-crisis market power is associated to a receivables days period that is approximately 3 days longer, confirming Meltzer (1960), Biais and Gollier (1997), Wilner (2000), Cuñat (2007), Petersen and Rajan (1997) and Fabbri and Klapper (2016). This effect however, is not materially different during the crisis (i.e., the effect of a one standard deviation change in market power on receivables is still 3 days). Therefore, firms with high market power increase their net trade credit days during the crisis (relative to low market power firms), mainly by paying their suppliers earlier, and not by extending more trade credit to their customers.

We are cautious about interpreting our results as being the causal effect of market power on trade credit, because the assumptions needed for such an interpretation of our empirical results are obviously not testable. A causal interpretation of our results requires the cross sectional variation in pre-crisis market power to be unrelated to the change in the firms features that drive trade credit decisions, at both the supply and the demand levels. To reduce concerns about possible biases related to this issue, we do a series of exercises to improve our identification, and several robustness checks to consider possible alternative explanations and confounding effects.

One possible (and important) confounding effect is that financial constraints could be related to firm market power. A large strand of the literature predicts that the demand for trade credit increases during periods of monetary tightening [Biais & Gollier (1997)]. Empirical evidence shows that trade credit can indeed redistribute liquidity from firms with access to financial markets to constrained firms during such periods [Meltzer (1960), Schwartz (1974), Fazzari, Hubbard, and Petersen (2000), Nilsen (2002), Fisman and Love (2003)], offsetting, at least in part, the bank lending channel effect of scarce liquidity. However, recent empirical studies [Love, Preve, and Sarria-Allende (2007), Love and Zaidi (2010), Garcia-Appendini and Montoriol-Garriga (2013)] show that when a drastic reduction in the external supply of funds occurs, such as during a financial crisis, interfirm trade credit redistributes liquidity in a limited manner. As the crisis widens and deepens, even liquidity-rich firms may lose their ability to access bank and capital markets and cut trade credit as a result.

Indeed, the study of Garcia-Appendini and Montoriol-Garriga (2013) uses the 2007/08 credit crisis to identify the effect of financial constraints in trade credit extension, and finds that financially constrained firms reduce the supply of trade credit during the crisis. Therefore, we do a number of tests to assure that our results are not driven by some correlation between market power and financial constraints. These robustness tests aim to disentangle the different roles of market power and of financial constraints on trade credit. First, we directly control for the effects of financial constraints on trade credit, using control variables that account for financial distress (cash holdings, short-term debt, cash flow, firm size, fixed capital and free collateral). Our results are compliant with the findings of Garcia-Appendini and Montoriol-Garriga (2013), in that we show that pre-crisis liquidity is an important determinant of trade credit. Notwithstanding, the effect of market power remains virtually unchanged. We run separate regressions for financially constrained and unconstrained firms and market power appears as a determinant of trade credit for both groups, therefore eliminating the possibility of a confounding effect that would explain our results. We also present a series of other pieces of evidence, discussed in the fourth section of the paper, to assure that our results are not simply driven by financial constraints.

We also account for the differential effects of the crisis on the supply and demand for trade credit at the industry level by using industry-quarter fixed effects. Finally, our results hold when we use other measures of market power, at both the firm and the industry level. For example, we use the US Census data to provide a structural measure of market power at the industry level, and we conclude that firms in more concentrated sectors (i.e., more likely to be market powerful) reduce payable days during the crisis in comparison to firms in less concentrated industries.

We go further and analyze other variables after the crisis to understand the mechanism driving the increase in net trade credit days for high market power firms. We find that these firms increase sales relative to low market power firms during the crisis, which is consistent with an effort by these firms to maintain their monopoly rents.

This paper contributes to the previous literature by considering the impact of competition in the product markets and bargaining power as important decision factors for the company to extend trade credit. In particular, we add evidence that trade credit can be used for strategic purposes in the product market, which is consistent with some of the market power theory papers. We also add evidence on the mechanism driving trade credit decisions and their impacts on costs and increase in sales. Our results add to Fabbri and Klapper (2016), by confirming that financial constraints are important drivers of trade credit decisions. They also add to Dass, Kale, and Nanda (2015) by providing evidence on the importance of firm market power to trade credit decisions, according to their predictions, and adding a cleaner identification strategy that relies on a source of variation (the financial crisis) that is less endogenous to firms' decisions.

More importantly, this paper adds to a growing body of literature on trade credit during financial crises. Most of the existing literature addresses the role of financial constraints and the redistribution theory during a crisis [Love, Preve, and Sarria-Allende (2007), Love and Zaidi (2010), Kestens, Cauwenberge, and Bauwhede (2012), Garcia-Appendini and Montoriol-Garriga (2013), Casey and O'Toole (2014), Carbó-Valverde, Rodríguez-Fernández, and Udell (2016)], whereas this paper adds another important dimension to trade credit decisions during a crisis. Our results imply that firms with similar access to liquidity but a different degree of market power respond very differently to a credit supply shock. We contribute to the literature concerning trade credit by showing that the effects reported by the previous literature are better explained by including market power as an important additional factor. Results reported in other papers [e.g., Love, Preve, and Sarria-Allende (2007) and Murfin and Njoroge (2015)] regarding the decrease in the redistribution role of trade credit during crises may be mainly driven by firms operating in highly competitive markets. To the best of our knowledge, no paper explores credit crises to study market power as an important trade credit driver.

The remainder of the paper proceeds as follows. Section 2 describes the data. Section 3 introduces the empirical strategy. Section 4 presents the results and provides some robustness checks, and Section 5 concludes the paper.

## 2. DATA

Our sample is drawn from Compustat North America Fundamentals Quarterly. We collect data for U.S. companies from 2004 to 2010 to have a symmetrical panel around the crisis onset. We follow Duchin, Ozbas, and Sensoy (2010) and define the beginning of the credit crisis as the third quarter of 2007. To avoid survival bias, we collect all quarterly data available for each firm. We exclude firms in the financial sector, real estate, social services and public administration industries (NAICS 52, 53, 83 and 92). We also drop observations with receivable or payable days above 365 days or below zero (following Love et al. (2007)) or with gross margin below -100% (because they are likely to be measurement errors). A detailed description of the sample selection filters for 2007Q2 is provided in appendix 1. Our sample consists of 8,602 firms and a total of 123,552 observations.

The variables used in our main specifications and robustness checks are defined in detail in Appendix 2. Accounts receivable and accounts payable are scaled by the effective daily sales and cost of goods sold for each quarter, respectively, yielding the variables *Receivable Days* and *Payable Days*, which may be interpreted as the average number of days to receive sales and to pay suppliers. They are obtained by the ratio of receivables and payables to daily sales and the daily cost of goods sold, respectively. The daily sales and cost of goods sold are calculated by dividing quarterly sales and cost of goods sold (COGS) by the actual number of days in each quarter. This procedure follows similar studies (e.g., Love et al., 2007). Alternatively, we also scale account receivables and payable to total assets. All variables in our models were winsorized at the $2^{nd}$ and $98^{th}$ percentiles to mitigate the effect of outliers. Because working capital accounts are jointly determined by firms' policies, which in many cases try to match maturities or cash flows, we follow Love et al. (2007) and include the dependent variable

*Net Trade Credit Days* to analyze the decision to increase marginal trade credit extension during the crisis.

We also analyze firms' *Sales to Assets* ratio (turnover), *Payables to Assets*, *Receivables to Assets* and *Gross Margin* as dependent variables to understand the different policies adopted by the firms with respect to trade credit extension relative to their total assets.

Our controls are the commonly used liquidity indicators, such as cash and equivalents, operational cash flow and short-term debt, all of them scaled by total assets, following Love, Preve, and Sarria-Allende (2007) and Duchin, Ozbas, and Sensoy (2010). We also control for Property, Plant and Equipment (PPE) to Assets, as firms with more collateral may follow distinct trade credit policies, as in Garcia-Appendini and Montoriol-Garriga (2013). We also use a proxy for free collateral, defined as (PPE – Long Term Debt) / Assets and relationship-specific investment (RSI) as in Dass, Kale, and Nanda (2015).

As explained before, one possible confounder in our analysis could be the firms' level of financial constraints. To address this problem, we run separate regressions for constrained and unconstrained firms. We measure financial constraints by the Whited-Wu index, the Hadlock-Pierce size-age (S.A.) index and firms' size measured by their natural log of quarterly sales [Whited and Wu (2006), Hadlock and Pierce (2010), Fabbri and Klapper (2016), Love and Zaidi (2010), Atanasova (2007)]. Detailed formulas for these indexes are presented in appendix 2 and follow the original definitions presented by their authors.

## 3. EMPIRICAL STRATEGY

To examine the extent to which market power affects trade credit, we exploit the variation in the importance of market power to trade credit decisions, as explained earlier. We estimate the following model to analyze the differential impact of market power in the trade credit variables:

$$TCvar_{i,j,t} = \sum_{T=1}^{3} \omega_T \times MP_i \times Post_T + \beta' \times Control_i \times Post_T + \mu_i + YQ_t + \varepsilon_{i,j,t}$$

(1)

where $TCvar_{ijt}$ is the trade credit dependent variable for firm $i$ in industry $j$ at quarter $t$.[4] We use three different dependent variables that capture distinct aspects of trade credit: *Receivable Days*, *Payable Days,* and *Net Trade Credit Days* in our main regressions.

We follow Dass, Kale, and Nanda (2015) and Gaspar and Massa (2006) and use the Lerner Index as the price-cost margin as follows:

$$Lerner\ Index_{i,t} = (Revenues_{i,t} - COGS_{i,t})/Revenues_{i,t}. \tag{2}$$

$MP_i$ is our main measure of market power, and is given by the quarterly average Lerner Index between 2004 and 2006 for firm $i$.[5] We use pre-crisis market power to avoid the possible effects of the shock on the variable, following the recommendation of Angrist and Pischke (2008). The main reason to use the 12-quarter average of the Lerner Index (instead of the value observed immediately prior to the crisis) is to mitigate any concerns on possible measurement error, as the Lerner Index could be influenced by short-term shocks faced by the firm (such as a disruption in production, for example), which are unrelated to actual market power. In robustness checks, we also use a dummy for firms with high market power, indicating that the firm market power ($MP_i$) is above median [similar to Fabbri and Klapper (2016)]. In further robustness checks, we use a measure of market power at the industry level (and therefore the variable becomes $MP_j$), taken from the US Census (see section 4 for detailed explanation).

$Post_T$ is a dummy set to one for the $T^{th}$ year after the crisis (so that $Post_1$ is a dummy for the first 4 quarters after the crisis, $Post_2$ is a dummy for the 5th to 8th quarters after the crisis and so on). Therefore, $Post_1$ equals 1 between 2007Q3 and 2008Q2 and 0 otherwise, $Post_2$ is equal to 1 between 2008Q3 and 2009Q2 and 0 otherwise, and $Post_3$ equals 1 between 2009Q3 and 2010Q2 and 0 otherwise. In some regressions, we collapse $Post_1$, $Post_2$ and $Post_3$ into a single post-crisis dummy ($Post_t$) to facilitate interpretation.

[4] The subscripts $i$ and $j$ are redundant in this case.

[5] In unreported robustness checks, we substitute the 2004-2006 average for the 2004 and 2006 averages, and all our inferences are maintained.

*$Control_i$* is a matrix containing firm-level control variables, also measured in the pre-crisis period (quarterly average from 2004 to 2006). They include liquidity and financial constraints indicators for firm *i*, namely, *cash / assets*, *Short-term Debt / Assets*, *Operational Cash Flow / Assets*, availability of collateral measures, given by *PPE / Assets* or *Free Collateral / Assets* . In some regressions, we also include investment in R&D as a proxy for *Relationship Specific Investment* (RSI), following the argument of Dass, Kale, and Nanda (2015) that trade credit may serve as a commitment device for a supplier to invest in a specific customer firm. Unobserved firm fixed effects that account for time-invariant firm-level unobserved heterogeneity are represented by $\mu_i$. Note that firm fixed effects subsume $MP_i$ and firm-level controls (not interacted with the $Post_T$ dummies). Time (year-quarter) fixed effects, $YQ_t$, capture macroeconomic fluctuations that homogeneously affect trade credit for all the firms in the sample (and subsume the $Post_T$ variables).

In alternative specifications, we fully saturate the model with industry-quarter fixed effects that account for industry-specific shocks that homogeneously affect all the firms in a given industry *j* in a given quarter *t* (e.g., industry seasonality). Industry-quarter fixed effects are also particularly relevant to mitigate concerns about changes in the demand for trade credit by downstream firms and the supply of trade credit by upstream firms during the crisis, that could be correlated to the firm's market power. Assuming that firms within an industry have similar suppliers and clients, industry-quarter fixed effects capture most of the variation in the supply and demand of trade credit, reducing concerns about possible omitted variable bias. Finally, $\varepsilon_{ijt}$ is the error term. Industry-quarter fixed effects also mitigate concerns about a possible measurement error in our measure of market power. The price-cost margin could be capturing abnormal profits that derive from short-term industry shocks unrelated to market power (for example, a temporary increase in the price of inputs), and therefore industry-quarter fixed effects are able to absorb these short-term fluctuations that occur at the industry level. Another reason to include industry-quarter fixed effects is to mitigate measurement error in the sense that competitive, but fast growing industries could present higher Lerner Indices, which would be a source of measurement error to our main measure of market power.

The parameters to be estimated are the vectors *β'* and *ω'*. Our main parameters of interest are the coefficients in the vector *ω'*, which represent the differential effect of market power (i.e. the effect of market power during the crisis as compared to normal times) on the dependent variables of trade credit.

We estimate Equation (1) with several variations. We start with a pooled OLS (i.e., without any fixed effects) with no controls, such that we can include $MP_i$ and the three $Post_T$ dummies, (without any interactions). We follow by introducing firm and time fixed effects, and therefore omit non-interacted pre-crisis $MP_i$ and controls and $Post_T$ dummies, because they are respectively firm-invariant time-invariant, and therefore are subsumed by their respective fixed effects. Finally, we saturate the model with industry-quarter fixed effects instead of time fixed effects. In some estimations, we collapse *Post1*, $Post_2$ and $Post_3$ into a single post-crisis dummy to facilitate interpretation. Standard errors are clustered at the NAICS 6-digit industry level to account for within-industry and time-series correlation of the residuals, following Bertrand, Duflo, and Mullainathan (2004).

# 4. DESCRIPTIVE STATISTICS AND REGRESSION RESULTS

## 4.1. Descriptive statistics

The results in table 1 show that the median firm extends approximately 50 days to its customers and pays its suppliers in 45 days. The *net trade credit days* variable has a negative mean of approximately 9 days. For comparison, we split the sample into firms with above and below median pre-crisis market power (*HiMP* and *LoMP* in table 1). High market power firms have larger receivable and payable periods, both before and after the onset of the crisis. The mean and median receivables period changes only slightly for both groups of firms from the pre- to the post-crisis period (and the pre- vs. post- differences for each group are not statistically significant). The mean payables period, however, is reduced from approximately 74 to 70 days for high market power firms, whereas it increases from approximately 50 to 54 days for low market power firms (differences are statistically significant at the 1% level). The average firm uses 21% of its sales as net trade credit. A median firm will show in its balance sheet 12.2% of its assets as receivables and 6.7% as payables.

Panel B of table 1 shows that the average gross margin (our measure of market power) is 37.3% for the whole sample, and the mean and median gross margins change only slightly from the pre- to the post-crisis period (change is not statistically significant), suggesting that firm market power is not severely affected by the crisis. The median firm keeps 10.5% of its assets as cash and equivalents, but firms in the 90$^{th}$ percentile hold as much as 53.3% in cash (not reported in table 1).

**[Insert Table 1 about here]**

## 4.2. Regression Results

### 4.2.1. Implications of Market Power for Trade Credit

The results for the estimation of our baseline model using net trade credit (*NetTC*) as the dependent variable are shown in Table 2. The estimations shown in Columns 1 to 4 use pooled OLS. Estimations using firm and time fixed effects are displayed in columns 5 to 7, and the estimation in column 8 adds industry-time (*industry* × *YQ*) fixed effects.

The results in column 1 of table 2 show that there is an average increase of approximately 5 days in net trade credit in the year that follows the crisis and approximately 3 and 4 days in the following years compared to other (normal) periods. In normal times, firms with greater market power have smaller net trade credit periods on average. An increase of one standard deviation in firm market power (22.7 percentage points) decreases net trade credit by approximately 12.8 days.[6] During the first year of crisis, however, this effect is reduced to 8.8 days, and in the second and third years of crisis the effect is approximately 9.5 days. All

[6] The computations are as follows. In normal times, an increase of one standard deviation in market power (22.7 percentage points) is *0.227 * (-56.342) ≈ -12.8* days, whereas during the first, second and third years of crisis, the effect is 0.227 * (-56.342 + 17.237) ≈ *-8.8* days. The effects for the second and third year of crisis are computed analogously.

coefficients are statistically significant at the usual levels. Therefore, high market power firms reduce their net trade credit days compared to low market power firms during the crisis.

**[Insert table 2 about here]**

One might suspect that liquidity issues, and not market power, drive these results, because firms with more market power could be holding more cash before the crisis. We test this possibility in the estimations in column 2 of table 2, by introducing pre-crisis cash positions and other firm-level controls for financial constraints (namely, cash flow, short-term debt) and potential collateral (*PPE / assets*) and their interactions with the crisis dummy (*Post*), following the arguments by Duchin, Ozbas, and Sensoy (2010) and Garcia-Appendini and Montoriol-Garriga (2013). In this regression, the signs and significance are maintained, but the coefficients of interest are increased in magnitude. According to this estimation, a one standard deviation increase in market power is associated with a decrease of almost 18 days in net trade credit, but the effect is reduced by approximately 5 days during the crisis. In column 3 of Table 2, we replace *PPE / Assets* with *Free Collateral / Assets* as our measure of available collateral, and the coefficients of interest change minimally. In column 4 of Table 2, we follow Dass, Kale, and Nanda (2015) and replace *PPE/Assets* with a control for relationship-specific investments (RSI), proxied by a dummy that captures above-average investment in R&D.[7] We find that relationship-specific investment increases net trade credit in normal times, consistent with Dass, Kale, and Nanda (2015), but this effect is reduced during the crisis.

The estimations in columns 5 to 7 of table 2 are respectively analogous to columns 2 to 4 respectively, but they introduce firm and time fixed effects. The results shown in these columns indicate that, in comparison to normal times, a one standard deviation increase in market power is associated with an increase in net trade credit of approximately 3 days in the year following that crisis and 4 days in the two subsequent years.[8] Importantly, once we introduce firm and

---

[7] We follow Dass, Kale, and Nanda (2015)main approach, in which uninformed investment in R&D is considered zero. In unreported robustness checks, we exclude observations in which R&D investment is uninformed, and we obtain qualitatively similar results.

[8] Taking the coefficients of column (5) as an example, the computation of the effect in the first year is given by *0.227*15.267 ≈ 3.5* days. In the second year, the effect is *0.227*16.783 ≈ 3.8* days, and in the third year after the crisis, the effect is *0.227*18.481 ≈ 4.2* days. All the effects estimates provided hereafter follow the same logic.

time fixed effects, pre-crisis *cash*, *cash flow* and *short term debt* do not seem to affect *net trade credit* in significantly different manner from what they do in normal times.[9] PPE / Assets is positively related to *net trade credit* (column 5), but when we replace it with a measure of *free collateral* (column 6) we do not find statistical significance.

Finally, by introducing industry-time fixed effects into the estimation shown in column 8 of table 2, we obtain very similar results, which strongly suggests that our results are not driven by any industry-related effect or industry-specific seasonality. To sum up, the coefficients measuring the differential effect of market power on net trade credit during the crisis reported in table 2 are significant both statistically and economically and are remarkably stable to several different specifications. These results reveal that a one standard deviation increase in market power is associated to an increase of 3-6 days in net trade credit during the crisis compared to normal times.

**[Insert table 3 about here]**

The estimations in Table 3 explore whether the changes in net trade credit days (analyzed in table 2) stem from variations in receivable and/or payable days. The results in column 1 of Table 3 show the results of pooled OLS regressions of *payable days* on *market power*, the 3 crisis dummies ($Post_1$, $Post_2$ and $Post_3$) and their interactions, without any control variables. Column 1 of table 3 shows that a one standard deviation increase in market power is associated to an increase of approximately 16 days in payables in normal times. This result is consistent with Long et al's. (1993) and Giannetti, Burkart, and Ellingsen's (2011) theories of product quality asymmetry by low market power firms. However, during the crisis, this effect is reduced. The coefficients of the crisis dummies indicate that payables increase by 5 days in the first year of crisis, and 3 days in the two following years, compared to normal times.

The regressions reported in columns 2 and 3 of table 3 include firm and time fixed effects, as well as controls for liquidity, and different measures of collateral availability. The

[9] In fact, the coefficient of *cash* is statistically significant at the 10% level in column 6, but the economic magnitude is small. A one standard deviation change in *cash* is associated to a change of less than 1 day in *net trade credit*.

coefficients of interest in these regressions differ only slightly from those reported in table 1, as they imply that a one standard deviation increase in market power is associated to a decrease of approximately 3, 4 and 5 days decrease in payables respectively in the first, second and third year of crisis compared to normal times. Finally, by introducing industry-time fixed effects in the regressions reported in column 4 of table 3, the coefficients $\omega$ are practically unchanged relative to the regressions reported in columns 2 and 3.

The fact that firms with high market power decrease net trade credit days more sharply in the years following the credit crunch supports the market power theory as found in Meltzer (1960), Biais and Gollier (1997), Wilner (2000), Cuñat (2007) and Petersen and Rajan (1997). Indeed, reducing payable days to support potentially financially constrained suppliers may help ensure a steady supply, and reduce concerns that suppliers will reduce the quality of products and services provided, which would prevent these high market power firms from losing their monopoly rents.

The estimations shown in columns 5 to 8 of table 3 are analogous to columns 1 to 4, but using *receivable days* as the dependent variable. The results of column 5 show that a one standard deviation increase in market power is associated to an increase of $0.227*13.724 \approx 3$ days in receivables, but we do not observe any difference during the crisis, as the three $\omega$ coefficients (of $MP_i \times Post_T$) are not statistically significant.

By introducing control variables and firm and time fixed effects in the regressions reported in columns 6 and 7 of column 3, we do not observe any material change to our inference from the OLS regression. The coefficients of interest for the first two years of crisis, $\omega_1$ and $\omega_2$ are not statistically significant. Despite the statistical significance of $\omega_3$ in these regressions, the correspondent magnitude of the economic effect is very small, implying that a one standard deviation increase in market power is associated to a decrease of less than 1 day in receivables during the crisis in comparison to normal times. Finally, we introduce industry-quarter fixed effects in the regression reported in column 8 of Table 3, and our coefficients of interest are not statistically significant. To sum up, the results reported in columns 5 to 8 of table 3 imply that *market power* is positively associated to *receivables days* in normal times, and this relationship does not change during the crisis.

Taken together, the estimations shown in table 3 show that the increase in net trade credit by higher market power firms during the crisis (observed from the results in table 2) is mainly due to a decrease in payable days by these firms in comparison to low market power firms. Therefore, our data show that firms that operate in low-competition markets (high market power) are able to provide more liquidity to their suppliers and probably take advantage of the early payment discount. These results are compliant with those of Klapper, Laeven, and Rajan (2012) and Giannetti, Burkart, and Ellingsen (2011). Because firms' cost of capital increases during the financial crisis, the embedded interest rates in the trade credit discount for two-part contracts become significantly larger than in normal times. Thus, high market power firms will pay their suppliers faster [Barbosa, Moreira, and Novaes (2017)].

The inclusion of several control variables mitigates any concerns about potential confounding effects. Importantly, our inferences about the influence of market power on trade credit seems to be independent of the effects of financial constraints, as the inclusion of these controls does not change our inferences significantly. Consistent with most of the literature, higher pre-crisis cash balances are associated with an increase in trade credit provided to clients (*receivable days*) in the three years of crisis. Taking the coefficients from column 8 as our preferred specification, the estimations indicate that a one standard deviation increase in pre-crisis cash holdings increases *receivable days* by approximately one day in the three years of crisis.[10] Our results also show that firms with higher pre-crisis cash holdings increase their *payable days* during the crisis by 1-2 days in the years that follow the crisis. We also find evidence that larger pre-crisis operational cash flows are associated with an increase in both payable days and receivables days after the crisis. Short-term debt is associated with an increase in receivables days, but not with payables days. Finally, the results from columns 2, 4, 6 and 8 of table 3 are also consistent with those of Dass  Kale, and Nanda (2015), in that they conclude

[10] The standard deviation of pre-crisis cash holdings is 0.193. Therefore, the increase in receivable days associated with a 1 standard deviation increase in cash holdings is $0.193 \times 6.987 \approx 1.3$ days in the first year after the crisis. Analogous calculations lead to an effect of 1.3 and 1.1 days in the second and third years of crisis respectively.

that relationship-specific investment (RSI) is important in explaining the supply of trade credit (i.e., *receivable days*) but not its uptake (i.e., *payable days*).

#### 4.2.1. Disentangling Financial Constraints

The estimations shown in tables 2 and 3 control for financial constraints using variables typically used in the literature for this purpose (such as pre-crisis cash holdings, cash flows and short-term indebtedness) as well as their interactions with the post-crisis dummies. However, one could still be concerned about a possible confounding effect. If firms with higher market power are also less financially constrained, the results shown in tables 2 and 3 could be driven by firms' financial constraints instead of market power. To further disentangle the effects of market power and financial constraints on trade credit during the crisis, we run separate regressions for subsamples of constrained and unconstrained firms, following Duchin, Ozbas, and Sensoy (2010) and Garcia-Appendini and Montoriol-Garriga (2013).

We run the baseline model without control variables (i.e., containing only the three interaction terms $MP \ x \ Post_T$ and firm and year-quarter fixed effects) specification for subsamples of financially constrained and unconstrained firms. We define these subsamples using three different measures of financial constraints: the Whited-Wu index [Whited and Wu (2006)], the Hadlock-Pierce SA index [Hadlock and Pierce (2010)], whose formulas are defined in Appendix 2. After computing each of these indicators at the firm level in the pre-crisis period, we split the sample into high constraints (above median) and low constraints (below median). The estimations reported in columns 1 and 2 of table 4 use firms with an above-median or below-median Whited-Wu index, respectively. Analogously, columns 3 and 4 of Table 4 use firms with an above-median or below-median Hadlock-Pierce SA index, respectively. Finally, we use a simpler definition of financial constraints in the estimations reported in columns 5 and 6, where we split the sample between below-median and above-median log of sales (proxy for size).[11]

[11] We use average 2004-2006 values of each of these indicators to split firms into the subsample of firms.

**[Insert table 4 about here]**

The regressions reported in table 4 use *Net Trade Credit* as the dependent variable. We verify that all the estimated $\omega$ coefficients are positive and statistically significant at the usual levels, indicating that our previous inferences about the role of market power on trade credit during a credit crisis hold for both constrained and unconstrained firms. A one standard deviation increase in market power is associated to a predicted increase in net trade credit during the crisis that ranges between 2 and 5 days, depending on the year and subsample considered. Although the magnitude of the coefficients differ across subsamples, we verify very stable positive and significant ω coefficients obtained for all models, for both constrained and unconstrained firms, regardless of the definition of financial constraints used.

Overall, the results in table 4 confirm that market power drives firms' trade credit decisions irrespective of their degree of financial constraints and that it is neither liquidity nor financial constraints that produce our main finding. Firms with high market power are more able than their low market power counterparts to keep the redistribution channel described in Meltzer (1960) and Cuñat (2007) working even during the crisis, particularly by providing liquidity to their suppliers.

Finally, to further eliminate any concerns that our measure of market power is essentially capturing financial constraints, we perform a two-stage procedure. First, we regress our market power variable on the three aforementioned proxies for financial constraints (Whited-Wu index, Hadlock-Pierce index and log of sales) and then use the residuals of these regressions (i.e., the portion of market power unexplained by financial constraints) in place of our original market power measure in our main specification.

**[Insert table 5 about here]**

The regression results of these exercises are reported in Table 5. Overall, our previous inferences are all sustained, including the magnitude of coefficients and statistical significance. These results strongly suggest that our findings that high market power firms reduce payable days during the crisis compared to their low market power counterparts does not stem from a possible confounding effect with financial constraints. The regression results are also consistent with our previous inference regarding *receivable days*. In the regressions reported in column 7 to 9 of table 6, we do not obtain statistical significance for the coefficients of the first two interaction terms ($\omega_1$ and $\omega_2$). The coefficient of the third interaction term $\omega_3$ is statistically significant, but has a small magnitude (a one standard deviation change in market power is associated to a change of less than one day in receivables).

### 4.2.2. Other robustness checks

In this section, we start by verifying whether our previous inferences are driven by a measurement error of market power. With this purpose, we use two alternative measures of market power. First, we modify our continuous measure of market power into a dummy variable ($HiMP_i$) that assumes value 1 if $MP_i$ is above the sample median and 0 otherwise. Second, we use a structural measure of market power at the industry level. We follow Ali, Klasa, and Yeung (2009) and gather data on market concentration from the US Census of 2007.[12] Specifically, we use the market share of the top 4 firms within an industry (4-digit NAICS) to measure industry concentration. We then build a dummy of highly concentrated industry ($HiMP_j$), that assumes value 1 for industries with concentration above the median, and 0 otherwise.

We report the regression results for these exercises in Table 6. To facilitate the interpretation, in all these regressions we collapse the three $Post_T$ dummies into a single dummy *Post* (such that *Post* assumes value 1 for the 12 quarters of crisis, and 0 otherwise). Column 1 of Table reports the regression results using *Net Trade Credit* as the dependent variable and the high market power dummy (*HiMPi*) as our measure of market power. The results indicate that high market power firms increase net trade credit during the crisis by approximately 4.8 days during the crisis in comparison to low market power firms, consistent with our previous evidence (statistically significant at the 1% level). The magnitude and significance of the effect is comparable to the ones obtained in Table 2, using our continuous measure of market power.

[12] As Ali et al. (2009) note, calculating market shares at the firm level using Compustat data can lead to severe measurement error, because the computation leaves out unlisted firms that may detain important shares of the market.

**[Insert Table 6 about here]**

In the regression reported in column 2 of Table 6, we use the industry-level measure of market power, and find that firms in less competitive industries increase net trade credit by 1.8 days more than firms in less competitive industries during the crisis (statistically significant at 5%). The fact that we find a smaller effect (1.8 days) using the industry-level variable compared to the firm-level variable (4.8 days) is expected, since the industry-level measure captures only the between-industry variation in market power, whereas the firm-level measure captures both between and within-industry variation in market power.

The regressions reported in columns 3 and 4 of Table 6 are analogous to the ones in columns 1 and 2 respectively, but using *payable days* as the dependent variable. These results confirm that most of the effect observed for *Net Trade Credit* stem from variations in *payable days*. The coefficient in column 3 shows that high market power firms decrease payables by 4.7 days compared to low market power firms during the crisis, whereas the results in column 4 indicates that firms in low competition industries decrease payables by 1.6 days compared to firms in high competition industries during the crisis. Finally, the regressions reported in columns 5 and 6 of table 6 use *receivable days* as the dependent variable, and the coefficients confirm that there is little additional effect of market power on *receivable days* during the crisis.

Overall, the results reported in Table 6 confirm our previous inferences. Importantly, the results using a structural measure of market power (industry concentration) are very compelling because they confirm inferences made from tests using a very different, non-structural measure of market power (the Lerner index), and these variables are computed in very different manners using different sources of data.

One might also suspect that firms with greater market power could anticipate their change in trade credit policy, as the credit shock precedes the demand shock, and thus, our results could be driven by the demand shock that started in September 2008 and not by the credit crisis. To validate our identification strategy, we follow the idea of Duchin, Ozbas, and Sensoy (2010) and Almeida et al. (2012) and run regressions with a placebo crisis in September 2004, when there was abundant liquidity and no signs of a credit crunch.

**[Insert Table 7 about here]**

Table 7 shows the results for the estimation of equation (1) using time and firm fixed effects for the placebo crisis in September 2004. Columns 1, 3 and 5 report the regression results respectively for *Net Trade Credit*, *Payable Days* and *Receivable Days* without any controls. In the regressions reported in columns 2, 4 and 6 of Table 7, we add controls to the previous regressions. Our results show that we cannot reject the hypothesis that the coefficients for the interaction of high market power and post-crisis year dummies are zero. Similar results are found for pooled regressions with and without control variables. Comparing the results in table 7 with those from the previous tables adds further evidence that our results are induced by the credit crisis.

Another possible concern of our model is that firm size (and not market power) affects trade credit during the crisis. We exclude this possibility by including pre-crisis firm size (measured by the average natural logarithm of quarterly sales between 2004 and 2006) in our regressions. The results (unreported) show that the inclusion of size in our regressions does not materially change our previous findings. A final possible cause of error in our models would be the possibility of correlation of the residuals between our clusters defined at the 6-digit NAICS level, causing standard errors to be underestimated. The estimations using clustering at the 4-digit NAICS level (unreported) is only slightly changed compared to the previous results.

### 4.2.3. Evidence on the mechanism

Because the crisis starts in 2007 with the bank credit reduction and higher interest rates, one should expect that firms would increase their price elasticity of demand for their inputs. As the crisis deepens after September 2008 and the recession hits the product market, we expect this effect to increase. Therefore, firms with higher market power could use their flexibility to decrease their higher margins to avoid losing sales or even increase their sales, as they acknowledge that their customers have increased their price elasticity. To analyze whether this mechanism could be in place for firms with higher market power during the crisis, we also estimate several alternatives to equation (1) using *Gross Margin, Sales / Assets*, *COGS / Assets* and *Inventory Days* as dependent variables.

**[Insert Table 8 about here]**

Column 1 of Table 8 shows that firms with higher market power decrease their gross margins, as we suspected. Indeed, a one standard deviation increase (decrease) in pre-crisis market power is associated to a predicted reduction (increase) of 0.3 percentage points in the after-crisis gross margin. Although small in magnitude, the effect is statistically significant at the 1% level. By reducing their margins, high market power firms are able to obtain an increase in their *sales / assets* ratio relative to low market power firms. The results reported in column 2 shows that a one standard deviation increase in market power is associated to an increase in *sales / assets* of 1 percentage point (statistically significant at the 1% level). Consistent with the increase in sales by high market power firms, the results in column 3 of table 8 show that a one standard deviation increase in market power is associated to an increase of approximately 0.2 percentage points. A similar effect is found for *payables / assets* in column 4. Note that this increase in *sales / assets* and *receivables / assets* for firms with higher market power is consistent with the insignificant effect of market power on *receivable days*, since it is plausible that high market power firms have increased sales by slightly reducing their margins, but without changing receivables policy.

Another possible mechanism underlying the previously reported reduction in *payable days* of high market power firms during the crisis is a larger mechanical reduction of *inventory days* for these firms. The results in column 5 of table 8 show that a one standard deviation increase in market power is associated to a decrease of approximately 2 days in inventories. This could explain only part (less than half) of the larger reduction in payable days of high market power firms and only in the first few months following the crisis. The relative reduction in inventory by high market power firms compared to their low market power counterparts also reinforces these firms' lower need for the financing of payables, as high market power firms are able to reduce inventories and still increase sales relative to low market power firms during the crisis.

## 5. Concluding remarks

This paper investigates whether product market power affects trade credit decisions, using a sample of Compustat U.S. firms from 2004 to 2010. We exploit the 2007-08 credit crisis in the U.S. as a source of variation in firm external financing conditions, which in turn affect the importance of financing customers and and liquidity-insuring suppliers. Previous works, such as Love, Preve, and Sarria-Allende (2007) and Garcia-Appendini and Montoriol-Garriga (2013), show that trade credit extension decreases during financial crises and that the redistribution theory does not hold during crises, because even firms with more access to the credit market would have lower liquidity available to be distributed to their customers. These works also find that constrained firms take more trade credit during the crisis and that firms with higher liquidity pre-crisis will extend more trade credit relative to ex ante low-cash firms.

Our paper confirms most of the previous findings about the effect of financial constraints on trade credit decisions during a crisis. More importantly, our results show that product market power is an important determinant of trade credit decisions, which has been overlooked by previous research in the field. We find that a one standard deviation increase in pre-crisis market power is associated to an increase in net trade credit of approximately 4-5 days during the crisis compared to normal times. We claim that this result is statistically and economically significant, representing approximately 10% of the median firm payable days in our sample. We conduct several tests, and our results are robust to all of them. In particular, we rule out the possibility that financial constraints could drive our findings by estimating separate regressions for constrained and unconstrained firms using the Whited-Wu index, the Hadlock-Pierce S.A. index, and the firms’ size, and we find that the effects of product market power on trade credit are present in both constrained and unconstrained firms. Therefore, financial constraints, liquidity and other control variables do not change our results. We also control for other measures of access to external funding, and use different measures of market power, structural and non-structural, and our results withstand all of these robustness checks. Finally, we use a placebo credit crisis and do not find any effects, as expected.

High market power firms are able to provide liquidity to their suppliers by decreasing their payable days, on average. This policy, besides possibly taking advantage of early payment discounts, allows these firms to inject extra liquidity into their suppliers during the credit crunch period, thus guaranteeing their inputs and the maintenance of monopoly rents. Since these firms operate with high price-cost margins, they also increase sales relative to firms with low market power by cutting their margins after the crisis onset. It is also possible that, due to the low liquidity in the credit market, suppliers of high market power firms find it difficult to obtain bank loans secured by their receivables and will thus increase the early payment discount to obtain liquidity from their customers.

This study contributes to the corporate financial policy literature by introducing an unexplored dimension to trade credit extension during a credit crisis. We show that trade credit extension and uptake by firms with similar liquidity but different market power may respond very differently to a credit supply shock. This paper is also related to the literature concerning investment in trade credit and financial constraints, and because it introduces a potentially omitted effect in previous works, it can enhance the explanation of the effects documented by Love, Preve, and Sarria-Allende (2007), Love & Zaidi (2010), Carbó-Valverde et al. (2016), Kestens et al. (2012) and Garcia-Appendini & Montoriol-Garriga (2013).

## 6. References


Ali, Ashiq, Sandy Klasa, and Eric Yeung, 2009, The limitations of industry concentration measures constructed with compustat data: Implications for finance research, *Review of Financial Studies* 22, 3839–3871.

Almeida, H, M Campello, B Laranjeira, and S Weisbenner, 2009, Corporate debt maturity and the real effects of the 2007 credit crisis *National Bureau of Economic Research*.

Angrist, JD, and JS Pischke, 2008, *Mostly Harmless Econometrics: An Empiricist's Companion*. Ed. Princeton University Press.

Atanasova, Christina, 2007, Access to Institutional Use Finance and the of Trade Credit, *Financial Management* 36, 49–67.

Barbosa, Klenio, Humberto Moreira, and Walter Novaes, 2017, Interest Rates in Trade Credit Markets, *Forthcoming, Journal of Money, Credit and Banking* 49, 75–113.

Bertrand, Marianne, Esther Duflo, and Sendhil Mullainathan, 2004, How Much Should We Trust Differences-in-Differences Estimates ?, *Quarterly Journal of Economics*, 249–275.

Biais, Bruno, and Christian Gollier, 1997, Trade Credit and Credit Rationing, *The Review of Financial Studies* 10, 903–937.

Carbó-Valverde, Santiago, Francisco Rodríguez-Fernández, and Gregory F. Udell, 2016, Trade Credit, the Financial Crisis, and SME Access to Finance, *Journal of Money, Credit and Banking* 48, 113–143.

Casey, E, and CM O'Toole, 2014, Bank lending constraints, trade credit and alternative financing during the financial crisis: Evidence from European SMEs, *Journal of Corporate Finance* 27, 173–193.

Collins, NR, and LE Preston, 1969, Price-cost margins and industry structure, *The Review of Economics and Statistics*.

Cuñat, Vicente, 2007, Trade credit: Suppliers as debt collectors and insurance providers, *Review of Financial Studies* 20, 491–527.

Daripa, A, and J Nilsen, 2011, Ensuring sales: A theory of inter-firm credit, *American Economic Association* 3, 245–279.

Dass, Nishant, Jayant R. Kale, and Vikram Nanda, 2015, Trade Credit, Relationship-specific Investment, and Product Market Power, *Review of Finance* 19, 1867–1923.

Duchin, R, O Ozbas, and BA Sensoy, 2010, Costly external finance, corporate investment, and the subprime mortgage credit crisis, *Journal of Financial Economics* 97, 418–435.

Fabbri, D, and L Klapper, 2016, Bargaining power and trade credit, *Journal of Corporate Finance* 41, 66–80.

Fabbri, D, and LF Klapper, 2008, Market power and the matching of trade credit terms *World Bank Policy Research Working Pape 4754*.

Fazzari, Steven M., R. Glenn Hubbard, and Bruce C. Petersen, 2000, Investment-cash flow sensitivities are useful: A comment on Kaplan and Zingales, *Quarterly Journal of Economics* 115, 695–705.

Fisman, Raymond, and Inessa Love, 2003, Trade Credit , Financial Intermediary and Industry Growth Development , *The journal of Finance* 58, 353–374.

Garcia-Appendini, E, and J Montoriol-Garriga, 2013, Firms as liquidity providers: Evidence from the 2007–2008 financial crisis, *Journal of Financial Economics* 109, 272–291.

Gaspar, JM, and M Massa, 2006, Idiosyncratic Volatility and Product Market Competition Published by : The University of Chicago Press Stable URL : http://www.jstor.org/stable/10.1086/505251 Idiosyncratic Volatility and Product Market Competition *, *the Journal of Business* 79, 3125–3152.

Giannetti, M., M. Burkart, and T. Ellingsen, 2011, What You Sell Is What You Lend? Explaining Trade Credit Contracts, *Review of Financial Studies* 24, 1261–1298.

Giannetti, Mariassunta, Mike Burkart, and Tore Ellingsen, 2011, What you sell is what you lend? Explaining trade credit contracts, *Review of Financial Studies* 24, 1261–1298.

Hadlock, CJ, and JR Pierce, 2010, New evidence on measuring financial constraints: Moving beyond the KZ index, *Review of Financial studies* 23, 1909–1940.

Kestens, Katrien, Philippe Van Cauwenberge, and Heidi Vander Bauwhede, 2012, Trade credit and company performance during the 2008 financial crisis, *Accounting & Finance AFAANZ* 54, 1125–1151.

Klapper, L, L Laeven, and R Rajan, 2012, Trade credit contracts, *Review of Financial Studies* 25, 838–867.

Long, Michael S, Ileen B Malitz, S Abraham Ravid, and S Long, 1993, Trade Credit, Quality Guarantees, and Product Marketability, *Financial Management* 22, 117–127.

Love, Inessa, Lorenzo A. Preve, and Virginia Sarria-Allende, 2007, Trade credit and bank credit: Evidence from recent financial crises, *Journal of Financial Economics* 83, 453–469.

Love, Inessa, and Rida Zaidi, 2010, Trade Credit, Bank Credit and Financial Crisis, *International Review of Finance* 10, 125–147.

Mateut, S, 2014, Reverse trade credit or default risk? Explaining the use of prepayments by firms, *Journal of Corporate Finance* 29, 303–326.

Meltzer, AH, 1960, Mercantile credit, monetary policy, and size of firms, *The Review of Economics and Statistics* XLII, 429–438.

Murfin, Justin, and Ken Njoroge, 2015, The implicit costs of trade credit borrowing by large firms, *Review of Financial Studies* 28, 112–145.

Nilsen, Jeffrey H, 2002, Trade Credit and the Bank Lending Channel, *Journal of Money, Credit and Banking* 34, 226–253.

Petersen, Michell A, and Raghuram G Rajan, 1997, Trade credit: theories and evidence, *The Review of Financial Studies* 10, 661–691.

Schmalensee, R, 1989, Inter-industry studies of structure and performance, *Handbook of industrial organization*.

Schwartz, RA, 1974, An economic model of trade credit, *Journal of Financial and Quantitative Analysis* 9, 643–657.

Silva, Dany Rogers, Kárem Cristina De Sousa Ribeiro, and Hsia Hua Sheng, 2011, Mensuração da rentabilidade do crédito comercial: aplicação em um caso atacadista-distribuidor. Vitória., *Brazilian Business Review* 8, 22–42.

Sutton, John, 2007, Chapter 35 Market Structure: Theory and Evidence, *Handbook of Industrial Organization* 3, 2301–2368.

Whited, TM, and G Wu, 2006, Financial constraints risk, *Review of Financial Studies* 19, 531–559.

Wilner, Benjamin S, 2000, The Exploitation of Relationships in Financial Distress : The Case of Trade Credit, *The Journal of Finance* 55, 153–178.

Wilson, Nicholas, and Barbara Summers, 2002, Trade Credit Terms Offered by Small Firms : Survey Evidence and Empirical Analysis, *Journal of Business Finance & Accounting* 29, 317–351.

***Table 1 - Descriptive Statistics***

This table reports summary statistics for our sample of firm-year-quarter observations from March 31, 2004, to June 30, 2010. All variables are defined in Appendix 2. The pre-crisis period is from 2004Q1 to 2007Q2, and the crisis period is from 2007Q3 to 2010 Q2. Panel A shows the descriptive statistics for trade credit measures, used as dependent variables in the main regressions. LoMP and HiMP firms are respectively defined as having below and above-median market power, as measured by the 2004-2006 average price-cost margin. Panel B shows the statistics for the measures of market power and liquidity, before and during the crisis. Panel C shows the descriptive statistics for the regression covariates as they are used in the regressions (12-quarter average from 2004 to 2006).

| | | Pre-Crisis | | | Crisis | | | Entire Period | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|
| Variable | | Mean | Median | Std. Dev. | Mean | Median | Std. Dev. | Mean | Median | Std. Dev. | Obs. |
| **A. Dependent Variables** | | | | | | | | | | | |
| | LoMP | 49.78 | 47.38 | 35.99 | 49.83 | 46.55 | 37.96 | 49.80 | 46.98 | 36.94 | |
| *Receivable Days* | HiMP | 54.98 | 53.35 | 33.09 | 55.26 | 53.00 | 34.84 | 55.10 | 53.24 | 33.86 | |
| | All firms | 52.36 | 50.46 | 34.68 | 52.29 | 49.51 | 36.68 | 52.33 | 50.03 | 35.60 | 123,552 |
| | LoMP | 49.89 | 38.89 | 45.82 | 54.11 | 39.89 | 53.51 | 51.90 | 39.37 | 49.68 | |
| *Payable Days* | HiMP | 74.30 | 53.14 | 66.35 | 69.74 | 50.16 | 62.88 | 72.32 | 51.77 | 64.90 | |
| | All firms | 62.01 | 44.93 | 58.23 | 61.20 | 44.30 | 58.46 | 61.64 | 44.62 | 58.34 | 123,552 |
| | LoMP | -0.11 | 4.07 | 48.45 | -4.28 | 2.59 | 56.01 | -2.10 | 3.32 | 52.23 | |
| *Net Trade Credit (days)* | HiMP | -19.32 | -4.07 | 70.45 | -14.48 | -2.14 | 67.75 | -17.22 | -3.27 | 69.33 | |
| | All firms | -9.64 | 0.85 | 61.14 | -8.91 | 0.98 | 61.82 | -9.31 | 0.91 | 61.45 | 123,552 |
| **B. Covariates** | | | | | | | | | | | |
| *Market Power (PCM)* | All firms | 37.3% | 35.9% | 24.8% | 37.2% | 35.7% | 24.9% | 37.3% | 35.8% | 24.8% | 123,552 |
| *Cash / Assets* | All firms | 18.8% | 10.5% | 20.6% | 18.2% | 10.6% | 19.8% | 18.5% | 10.5% | 20.2% | 123,445 |
| *Short-term Debt / Assets* | All firms | 5.3% | 1.0% | 10.0% | 5.5% | 1.1% | 10.4% | 5.4% | 1.0% | 10.2% | 120,091 |
| *PPE / Assets* | All firms | 25.6% | 17.2% | 23.6% | 26.7% | 17.6% | 24.7% | 26.1% | 17.4% | 24.1% | 123,377 |
| *Free Collateral* | All firms | 7.4% | 6.5% | 29.2% | 8.6% | 7.0% | 29.7% | 7.9% | 6.7% | 29.4% | 122,408 |
| *Oper. Cash Flow / Assets* | All firms | 1.7% | 2.8% | 6.6% | 1.7% | 2.6% | 6.1% | 1.7% | 2.7% | 6.4% | 116,634 |
| *Size* | All firms | 3.94 | 3.95 | 2.32 | 4.17 | 4.13 | 2.29 | 4.04 | 4.03 | 2.31 | 123,533 |
| | | | | | Mean | | Median | | Std. Dev. | | Obs. |
| **C. Covariate Pre-Crisis (average 2004 to 2006)** | | | | | | | | | | | |
| *MP (pre)* | | | | | 37.3% | | 35.3% | | 22.7% | | 117,837 |
| *Cash Avg (pre)* | | | | | 18.9% | | 11.5% | | 19.3% | | 117,828 |
| *Short-term Debt / Assets Avg (pr* | | | | | 5.2% | | 2.0% | | 8.5% | | 117,224 |
| *PPE / Assets Avg (pre)* | | | | | 25.9% | | 17.9% | | 23.2% | | 117,812 |
| *Cash Flow Avg (pre)* | | | | | 1.9% | | 2.9% | | 5.8% | | 114,487 |
| *Free Collateral (pre)* | | | | | 8.0% | | 6.6% | | 27.3% | | 117,754 |

**Table 2- Net Trade Credit Days after the 2007-8 Crisis**

This table presents estimates from pooled OLS and panel regressions explaining *Net Trade Credit* (days), which is measured by the difference between the number of days to receive sales (RecDays) and the number of days to pay suppliers (PayDays) for quarters ending between March 31, 2004, and June 30, 2010. We define post1, post2 and post3 as dummies for the first, second and third years after the onset of the crisis in the third quarter of 2007, respectively. We also define post as a dummy for the three years after the crisis start. MP is the ex-ante Lerner Index (PCM), measured as the average from 2004 to 2006. Cash is the average ex ante cash to assets ratio. Other covariates are defined in Appendix 2 and are measured prior to the crisis. Standard errors are robust to heteroscedasticity and are clustered at the industry level (6-digit NAICS). All model specifications employ robust standard errors, and robust t-statistics are reported in the parentheses below each coefficient. Superscripts ***, **, and * correspond to statistical significance at the one-, five-, and ten-percent levels, respectively. Constant and fixed effects coefficients are omitted.

| **Net Trade Credit (days)** | POLS | | | | Fixed Effects | | | |
|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| *Variables of interest* | | | | | | | | |
| *MP x post1* | 17.237 *** | 23.756 *** | 22.995 *** | 25.550 *** | 15.267 *** | 15.262 *** | 17.207 *** | 15.810 *** |
| | (4.829) | (7.239) | (6.533) | (7.003) | (5.920) | (5.686) | (6.140) | (4.828) |
| *MP x post2* | 14.604 *** | 23.098 *** | 22.159 *** | 25.010 *** | 16.783 *** | 16.791 *** | 18.702 *** | 18.359 *** |
| | (3.462) | (5.454) | (4.860) | (5.905) | (5.213) | (4.896) | (5.505) | (5.294) |
| *MP x post3* | 14.686 *** | 23.564 *** | 22.205 *** | 25.274 *** | 18.481 *** | 18.516 *** | 20.374 *** | 20.975 *** |
| | (3.508) | (5.744) | (5.252) | (6.364) | (5.493) | (5.197) | (5.689) | (5.504) |
| *post1* | -5.051 ** | -8.185 *** | -5.801 *** | -6.214 *** | | | | |
| | (3.229) | (5.112) | (3.465) | (3.388) | | | | |
| *post2* | -2.894 * | -7.555 *** | -5.072 ** | -5.617 *** | | | | |
| | (1.914) | (4.840) | (3.179) | (3.459) | | | | |
| *post3* | -3.766 ** | -8.566 *** | -6.035 *** | -6.917 *** | | | | |
| | (2.241) | (5.541) | (3.915) | (4.100) | | | | |
| *MP (Pre)* | -56.342 *** | -78.547 *** | -78.866 *** | -84.484 *** | | | | |
| | (4.186) | (6.942) | (5.642) | (4.723) | | | | |
| *Control variables (interacted)* | | | | | | | | |
| *Cash x post* | | -7.908 ** | -11.373 *** | -4.989 * | -2.150 | -4.320 * | -0.722 | -1.725 |
| | | (2.975) | (4.473) | (1.869) | (0.859) | (1.857) | (0.314) | (0.741) |
| *Short-term Debt x post* | | 14.753 * | 13.154 * | 12.812 | 7.181 | 6.389 | 6.262 | 6.237 |
| | | (1.830) | (1.650) | (1.615) | (1.078) | (0.954) | (0.936) | (0.937) |
| *Cash Flow x post* | | -55.866 *** | -47.644 ** | -66.140 *** | -2.940 | -0.436 | -12.171 | -6.877 |
| | | (3.968) | (3.127) | (4.208) | (0.235) | (0.035) | (0.849) | (0.527) |
| *PPE x post* | | 6.055 ** | | | 4.552 ** | | | 8.208 *** |
| | | (2.816) | | | (2.184) | | | (3.474) |
| *Free Collateral x post* | | | 0.933 | | | -0.086 | | |
| | | | (0.470) | | | (0.067) | | |
| *RSI x post* | | | | -134.454 *** | | | -108.411 ** | |
| | | | | (3.750) | | | (2.947) | |
| *Control variables (non-interacted)* | | | | | | | | |
| *Cash Avg (pre)* | | -11.891 ** | 21.724 ** | 10.510 | | | | |
| | | (2.360) | (2.025) | (1.003) | | | | |
| *Short-term Debt / Assets Avg (pre)* | | -63.295 *** | -54.670 *** | -50.765 *** | | | | |
| | | (7.933) | (6.257) | (5.472) | | | | |
| *Cash Flow Avg (pre)* | | 303.299 *** | 294.397 *** | 301.390 *** | | | | |
| | | (17.982) | (15.703) | (16.630) | | | | |
| *PPE / Assets Avg (pre)* | | -75.539 *** | | | | | | |
| | | (4.484) | | | | | | |
| *Free Collateral (pre)* | | | -41.685 ** | | | | | |
| | | | (2.912) | | | | | |
| *RSI (pre)* | | | | 368.957 ** | | | | |
| | | | | (3.238) | | | | |
| *Firm FE* | No | No | No | No | Yes | Yes | Yes | Yes |
| *Time FE* | No | No | No | No | Yes | Yes | Yes | No |
| *Industry x Time FE* | No | No | No | No | No | No | No | Yes |
| *Observations* | 117,837 | 113,951 | 113,945 | 113,951 | 113,951 | 113,945 | 113,951 | 113,951 |
| *R-squared* | 0.04 | 0.18 | 0.13 | 0.10 | 0.74 | 0.74 | 0.74 | 0.74 |
| *Clusters* | 887 | 870 | 869 | 870 | 870 | 869 | 870 | 870 |

**Table 3 - Impact of Market Power on Receivable and Payable Days**

This table presents estimates from pooled OLS and panel regressions explaining RecDays and PayDays, which are measured respectively by the number of days to receive from customers and to pay suppliers for quarters ending between March 31, 2004, and June 30, 2010. Net Trade Credit Days is the difference between these two variables, and therefore, they explain in detail the net effect in the previous table. We define post1, post2 and post3 as dummies for the first, second and third years after the onset of the crisis in the third quarter of 2007. We also define post as a dummy for the three years after the crisis start. MP is the ex-ante Lerner Index (PCM), measured as the average from 2004 to 2006. Cash is the average ex ante cash to assets ratio. Other covariates are defined in Appendix 2 and are always measured as the 2004-2006 (pre-crisis) average. Models are estimated using pool or unbalanced panel data with firm/time and industry/time fixed effects. Standard errors are robust to heteroscedasticity and are clustered at the industry level (6-digit NAICS). All model specifications employ robust standard errors, and robust t-statistics are reported in the parentheses below each coefficient. Superscripts ***, **, and * correspond to statistical significance at the one-, five-, and ten-percent levels, respectively. Constant and fixed effects coefficients are omitted.

| | Payable Days | | | | Receivable Days | | | |
|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| *Variables of interest* | | | | | | | | |
| *MP x post1* | -15.799 *** | -14.054 *** | -13.786 *** | -14.451 *** | 1.504 | 1.190 | 0.721 | 1.318 |
| | (5.376) | (6.179) | (6.174) | (4.543) | (0.958) | (1.007) | (0.608) | (1.201) |
| *MP x post2* | -15.075 *** | -18.364 *** | -18.105 *** | -18.483 *** | -0.661 | -0.724 | -1.199 | 0.494 |
| | (3.878) | (5.644) | (5.210) | (4.946) | (0.351) | (0.505) | (0.878) | (0.362) |
| *MP x post3* | -16.615 *** | -22.107 *** | -21.878 *** | -23.020 *** | -2.307 | -2.615 * | -3.076 ** | -1.242 |
| | (4.479) | (6.959) | (6.590) | (5.910) | (1.071) | (1.785) | (2.141) | (0.883) |
| *post1* | 5.372 *** | | | | 0.272 | | | |
| | (3.811) | | | | (0.421) | | | |
| *post2* | 3.384 ** | | | | 0.452 | | | |
| | (2.265) | | | | (0.582) | | | |
| *post3* | 3.378 ** | | | | -0.199 | | | |
| | (2.072) | | | | (0.242) | | | |
| *MP (Pre)* | 69.452 *** | | | | 13.724 *** | | | |
| | (5.499) | | | | (3.397) | | | |
| *Control variables (interacted)* | | | | | | | | |
| *Cash x post1* | | 4.691 * | 6.292 ** | 7.767 ** | | 6.658 *** | 5.450 ** | 6.987 *** |
| | | (1.725) | (2.403) | (2.872) | | (3.542) | (3.077) | (3.941) |
| *Cash x post2* | | 9.387 ** | 11.008 *** | 10.755 *** | | 7.250 *** | 6.050 *** | 6.551 *** |
| | | (2.886) | (3.814) | (3.796) | | (3.884) | (3.357) | (3.697) |
| *Cash x post3* | | 9.582 *** | 11.230 *** | 11.061 *** | | 5.989 ** | 4.776 ** | 5.559 ** |
| | | (3.478) | (4.408) | (3.939) | | (2.847) | (2.340) | (2.598) |
| *Short-term Debt x post* | | 0.364 | 0.764 | 1.113 | | 5.844 * | 6.123 * | 5.285 |
| | | (0.059) | (0.123) | (0.180) | | (1.765) | (1.866) | (1.588) |
| *Cash Flow x post* | | 41.294 *** | 38.129 ** | 40.940 ** | | 35.530 *** | 37.023 *** | 34.638 |
| | | (3.338) | (3.207) | (3.282) | | (5.211) | (5.390) | (5.052) |
| *PPE x post* | | -2.435 | | | | 0.903 | | |
| | | (1.008) | | | | (1.007) | | |
| *Free Collateral x post* | | | 1.154 | 0.145 | | | 0.880 | 0.315 |
| | | | (0.699) | (0.099) | | | (1.369) | (0.379) |
| *High RSI x post* | | 0.568 | | 0.286 | | -1.219 ** | | -1.819 ** |
| | | (0.615) | | (0.267) | | (2.634) | | (3.260) |
| *Firm FE* | No | Yes | Yes | Yes | No | Yes | Yes | Yes |
| *Time FE* | No | Yes | Yes | No | No | Yes | Yes | No |
| *Industry x Time FE* | No | No | No | Yes | No | No | No | Yes |
| *Observations* | 117,837 | 113,951 | 113,945 | 113,945 | 117,837 | 113,951 | 113,945 | 113,945 |
| *R-squared* | 0.07 | 0.72 | 0.72 | 0.73 | 0.01 | 0.79 | 0.79 | 0.79 |
| *Clusters* | 887 | 870 | 869 | 869 | 887 | 870 | 869 | 869 |

***Table 4- Impact of Financial Constraints in Trade Credit during the Crisis***

This table presents estimates from panel regressions explaining Net Trade Credit (days) for quarters ending between March 31, 2004, and June 30, 2010, for the whole sample (base model) as well as for subsamples including only firms with high (above median) or low (below median) financial constraints measured by the Whited-Wu index, the Hadlock-Pierce SA index and the size terms of the natural log of sales. Financial constraints measures are defined in Appendix 2 and are always measured as the 2004-2006 (pre-crisis) average. We define post1, post2 and post3 as dummies for the first, second and third years after the onset of the crisis in the third quarter of 2007, respectively. Market Power is the ex-ante Lerner Index (PCM) measured as the average from 2004 to 2006. Models are estimated using unbalanced panel data with firm and time fixed effects. All model specifications employ robust standard errors, and robust t-statistics are reported in the parentheses below each coefficient. Superscripts ***, **, and * correspond to statistical significance at the one-, five-, and ten-percent levels, respectively. Constant and fixed effects coefficients are omitted.

| **Net Trade Credit (days)** | Whited-Wu | | Hadlock-Pierce (SA) | | Size | |
|---|---|---|---|---|---|---|
| | High | Low | High | Low | Low | High |
| *MP x post1* | 16.305 *** | 11.587 *** | 16.037 *** | 9.634 ** | 16.557 *** | 8.740 ** |
| | (5.315) | (4.069) | (4.576) | (3.230) | (4.221) | (3.021) |
| *MP x post2* | 17.363 *** | 11.843 ** | 19.143 *** | 10.057 ** | 18.864 *** | 10.693 ** |
| | (3.697) | (3.075) | (4.324) | (2.487) | (3.609) | (2.901) |
| *MP x post3* | 22.087 *** | 12.300 ** | 21.844 *** | 10.459 ** | 24.075 *** | 7.348 ** |
| | (3.818) | (3.143) | (4.266) | (2.677) | (3.887) | (2.108) |
| *Firm FE* | Yes | Yes | Yes | Yes | Yes | Yes |
| *Time FE* | Yes | Yes | Yes | Yes | Yes | Yes |
| *Industry x Time FE* | No | No | No | No | No | No |
| *Observations* | 55,534 | 58,907 | 58,921 | 58,917 | 58,917 | 55,502 |
| *R-squared* | 0.79 | 0.70 | 0.80 | 0.81 | 0.70 | 0.68 |
| *Clusters* | 654 | 652 | 690 | 688 | 655 | 646 |

***Table 5 - Impact of Financial Constraints in Trade Credit during the Crisis***

This table presents estimates from panel regressions explaining Net Trade Credit (days), payable days (PayDays), and *receivable days* (RecDays) for quarters ending between March 31, 2004, and June 30, 2010. The original covariate Market Power (ex-ante Lerner index) is replaced by the residuals of regressions of this variable on the financial constraints indicators - firms with high (above median) or low (below median) financial constraints measured by the Whited-Wu index, the Hadlock-Pierce SA index and the size terms of the natural log of sales. We define post1, post2 and post3 as dummies for the first, second and third years after the onset of the crisis in the third quarter of 2007, respectively. We also define post as a dummy for the three years after the crisis start. Covariates are defined in Appendix 2 and are always measured as the 2004-2006 (pre-crisis) average. Models are estimated using unbalanced panel data with firm and time fixed effects. Standard errors are clustered at the industry level (6-digit NAICS). All model specifications employ robust standard errors, and robust t-statistics are reported in the parentheses below each coefficient. Superscripts ***, **, and * correspond to statistical significance at the one-, five-, and ten-percent levels, respectively. Constant and fixed effects coefficients are omitted.

| | Net Trade Credit | | | Payable Days | | | Receivable Days | | |
|---|---|---|---|---|---|---|---|---|---|
| | MP residual after WW | MP residual after SA | MP residual after Size | MP residual after WW | MP residual after SA | MP residual after Size | MP residual after WW | MP residual after SA | MP residual after Size |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
| *Variables of interest* | | | | | | | | | |
| *MP x post1* | 15.516 *** | 15.305 *** | 15.499 *** | -14.268 *** | -14.753 *** | -15.060 *** | 0.770 | 0.550 | 0.460 |
| | (6.121) | (5.855) | (5.726) | (5.861) | (6.406) | (6.350) | (0.711) | (0.518) | (0.428) |
| *MP x post2* | 16.682 *** | 16.975 *** | 17.432 *** | -18.184 *** | -18.021 *** | -18.319 *** | -1.761 | -1.170 | -1.013 |
| | (5.473) | (5.205) | (5.252) | (5.761) | (5.552) | (5.503) | (1.273) | (0.878) | (0.742) |
| *MP x post3* | 18.432 *** | 18.591 *** | 18.779 *** | -21.702 *** | -21.700 *** | -21.756 *** | -3.488 ** | -3.382 ** | -3.230 ** |
| | (5.884) | (5.480) | (5.475) | (6.696) | (6.706) | (6.607) | (2.411) | (2.392) | (2.254) |
| *Control variables (interacted)* | | | | | | | | | |
| *Cash x post* | -2.387 | -2.041 | -1.887 | 8.327 *** | 7.862 *** | 7.655 *** | 6.353 *** | 6.182 *** | 6.122 *** |
| | (0.968) | (0.817) | (0.760) | (3.802) | (3.691) | (3.623) | (3.647) | (3.689) | (3.668) |
| *Short-term Debt x post* | 6.805 | 7.600 | 7.878 | 0.612 | -0.239 | -0.536 | 6.208 * | 6.248 * | 6.231 * |
| | (0.991) | (1.140) | (1.181) | (0.095) | (0.039) | (0.087) | (1.789) | (1.888) | (1.884) |
| *Cash Flow x post* | -1.577 | -3.532 | -5.035 | 40.687 ** | 41.566 *** | 43.133 *** | 38.191 *** | 37.106 *** | 37.160 *** |
| | (0.125) | (0.282) | (0.401) | (3.229) | (3.481) | (3.608) | (5.694) | (5.387) | (5.378) |
| *PPE x post* | 4.562 ** | 4.535 ** | 4.565 ** | -2.276 | -2.759 | -2.798 | 2.039 ** | 1.606 * | 1.596 * |
| | (2.402) | (2.180) | (2.211) | (0.999) | (1.142) | (1.168) | (2.308) | (1.875) | (1.862) |
| *Firm FE* | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| *Time FE* | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| *Industry x Time FE* | No | No | No | No | No | No | No | No | No |
| *Observations* | 110,492 | 113,951 | 113,948 | 110,492 | 113,951 | 113,948 | 110,492 | 113,951 | 113,948 |
| *R-squared* | 0.74 | 0.74 | 0.74 | 0.72 | 0.72 | 0.72 | 0.79 | 0.79 | 0.79 |
| *Clusters* | 857 | 870 | 870 | 857 | 870 | 870 | 857 | 870 | 870 |

***Table 6 - Market Power impact on Trade Credit – Robustness Test***

This table presents estimates from panel regressions explaining Net Trade Credit (days), Payable Days and Receivable Days for quarters ending between March 31, 2004 and June 30, 2010. We define post as dummy for the first, second and third years after the onset of the crisis in the third quarter of 2007. High Market Power firms (*HiMP = 1*) have above median ex ante Lerner Index measured as the average from 2004 to 2006. High Market Power Census firms (*HiMP Census = 1*) have above median US Census Market Share of top 4 firms in 2007 (before crisis) for each NAICS 4-digit industry as in https://www.census.gov/econ/isp/. Cash is the average ex ante cash to assets ratio. Other covariates are defined in Appendix 2 and are always measured as the 2004-2006 (pre-crisis) average. Models are estimated using unbalanced panel data with firm/time fixed effects All model specifications employ robust standard errors, and robust t-statistics are reported in the parentheses below each coefficient. Superscripts ***, **, and * correspond to statistical significance at the one-, five-, and ten-percent levels, respectively. Constant and fixed effects coefficients are omitted.

| | Net Trade Credit | | Payable Days | | Receivable Days | |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| <u>*Variables of interest*</u> | | | | | | |
| *Hi MP x post* | 4.765 *** | | -4.725 *** | | -0.014 | |
| | (4.880) | | (4.904) | | (0.033) | |
| *Hi MP Census x post* | | 1.819 ** | | -1.630 * | | 0.220 |
| | | (2.158) | | (1.693) | | (0.459) |
| <u>*Control variables (interacted)*</u> | | | | | | |
| *Cash x post* | -1.622 | 1.439 | 6.258 ** | 3.162 | 5.062 ** | 4.974 ** |
| | (0.681) | (0.644) | (2.939) | (1.517) | (3.195) | (3.191) |
| *Short-term Debt x post* | 5.578 | 4.411 | 1.700 | 2.869 | 6.174 * | 6.193 * |
| | (0.838) | (0.660) | (0.274) | (0.457) | (1.877) | (1.873) |
| *Cash Flow x post* | 9.210 | 18.001 | 27.634 ** | 18.902 | 35.940 *** | 35.897 *** |
| | (0.738) | (1.407) | (2.329) | (1.547) | (5.423) | (5.315) |
| *PPE x post* | -0.127 | -0.156 | 1.197 | 1.245 | 0.894 | 0.918 |
| | (0.110) | (0.130) | (0.866) | (0.973) | (1.394) | (1.415) |
| *Firm FE* | Yes | Yes | Yes | Yes | Yes | Yes |
| *Time FE* | Yes | Yes | Yes | Yes | Yes | Yes |
| *Industry x Time FE* | No | No | No | No | No | No |
| *Observations* | 113,945 | 113,945 | 113,945 | 113,945 | 113,945 | 113,945 |
| *R-squared* | 0.74 | 0.74 | 0.72 | 0.72 | 0.79 | 0.79 |
| *Clusters* | 869 | 869 | 869 | 869 | 869 | 869 |

***Table 7 - Placebo Crisis and Trade Credit***

This table presents estimates from panel regressions explaining Net Trade Credit (days), Payable Days, Receivable Days, Payables to assets ratio and Receivables to assets ratio for quarters ending between March 31, 2003, and June 30, 2007. We define post as the dummy for the quarter after the placebo crisis on September/2004. Market Power is measured as the average Lerner Index before the placebo crisis. Covariates are defined in Appendix 2 and are always measured as the pre-crisis average. Models are estimated using unbalanced panel data with firm/time fixed effects. Standard errors are clustered at the industry level (6-digit NAICS). All model specifications employ robust standard errors, and robust t-statistics are reported in the parentheses below each coefficient. Superscripts ***, **, and * correspond to statistical significance at the one-, five-, and ten-percent levels, respectively. Constant and fixed effects coefficients are omitted.

| | Net Trade Credit | | Payable Days | | Receivable Days | |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| *MP x post* | 6.946 | 7.030 | -4.543 | -6.704 | 2.403 | 0.326 |
| | 1.546 | 1.569 | -1.050 | -1.630 | 1.288 | 0.270 |
| *Cash x post* | | 2.240 | | 3.547 | | 5.787 *** |
| | | 0.821 | | 1.370 | | 3.910 |
| *Short-term Debt x post* | | 10.401 * | | -11.893 ** | | -1.492 |
| | | 1.669 | | -2.135 | | -0.434 |
| *Cash Flow x post* | | -28.344 * | | 46.333 ** | | 17.989 ** |
| | | -1.906 | | 3.072 | | 3.101 |
| *PPE x post* | | -2.697 | | 3.927 | | 1.230 |
| | | -0.701 | | 0.999 | | 1.277 |
| *Firm FE* | Y | Y | Y | Y | Y | Y |
| *Time FE* | Y | Y | Y | Y | Y | Y |
| *Industry x Time FE* | N | N | N | N | N | N |
| *Observations* | 68,110 | 63,293 | 68,110 | 63,293 | 68,110 | 63,293 |
| *R2* | 0.77 | 0.77 | 0.75 | 0.75 | 0.79 | 0.80 |
| *Clusters* | 839 | 805 | 839 | 805 | 839 | 805 |

*Table 8- Market Power Mechanism during the Crisis*

This table presents estimates from panel regressions explaining Gross Margin, Sales to assets ratio, Receivables to assets ratio, and Inventory days for quarters ending between March 31, 2004, and June 30, 2010. We define post as dummy for the first, second and third years after the onset of the crisis in the third quarter of 2007. Market Power is the ex-ante Lerner Index measured as the average from 2004 to 2006. Cash is the average ex ante cash to assets ratio. Other covariates are defined in Appendix 2 and are always measured as the 2004-2006 (pre-crisis) average. Models are estimated using unbalanced panel data with firm/time fixed effects. Standard errors are clustered at the industry level (6-digit NAICS). Robust t-statistics are presented below each estimated parameter, and ***, ** and * represent the 1%, 5% and 10% significance levels, respectively.

| | Gross Margin | Sales to Assets | Receivables to Assets | Payables to Assets | Inventory Days |
|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) |
| *Variables of interest* | | | | | |
| *MP x post* | -0.135 *** | 0.046 *** | 0.011 ** | 0.011 ** | -9.669 *** |
| | (8.085) | (4.919) | (2.354) | (2.688) | (3.724) |
| *Control variables (interacted)* | | | | | |
| *Cash x post* | 0.047 *** | -0.038 *** | 0.004 | 0.003 | 11.274 *** |
| | (3.691) | (4.545) | (0.791) | (0.755) | (3.451) |
| *Short-term Debt x post* | -0.055 ** | -0.142 *** | -0.055 *** | -0.029 ** | 4.601 |
| | (2.583) | (4.921) | (4.163) | (2.305) | (0.849) |
| *Cash Flow x post* | -0.278 *** | -0.745 *** | -0.250 *** | -0.149 *** | 75.165 *** |
| | (5.465) | (12.276) | (11.816) | (6.131) | (5.821) |
| *Free Collateral x post* | -0.018 ** | 0.000 | 0.002 | 0.000 | 2.566 * |
| | (2.327) | (0.014) | (0.718) | (0.145) | (1.864) |
| *Firm FE* | Yes | Yes | Yes | Yes | Yes |
| *Time FE* | Yes | Yes | Yes | Yes | Yes |
| *Industry x Time FE* | No | No | No | No | No |
| *Observations* | 113,945 | 113,917 | 113,923 | 113,610 | 110,163 |
| *R-squared* | 0.82 | 0.83 | 0.86 | 0.83 | 0.85 |
| *Clusters* | 869 | 869 | 869 | 868 | 867 |

# Appendix I – Sample selection

***Table A1- Sample Selection from COMPUSTAT***

This table summarizes our sample selection filters applied to COMPUSTAT data to arrive at our final sample used in this paper. NAICS 52 correspond to financial firms and NAICS 53 corresponds to real estate and leasing firms.

| Period | **All firms** | minus NAICS 52 | minus NAICS 53 | minus Missing variables | minus Outliers | **Firms on Sample** |
|---|---|---|---|---|---|---|
| 2004-Q3 | **10,961** | *2,374* | *385* | *2,243* | *717* | **5,242** |
| 2004-Q4 | **10,802** | *2,388* | *379* | *1,529* | *780* | **5,726** |
| 2005-Q1 | **11,207** | *2,467* | *398* | *2,535* | *725* | **5,082** |
| 2005-Q2 | **11,116** | *2,477* | *393* | *2,335* | *730* | **5,181** |
| 2005-Q3 | **10,971** | *2,460* | *387* | *2,370* | *717* | **5,037** |
| 2005-Q4 | **10,775** | *2,467* | *382* | *1,590* | *779* | **5,557** |
| 2006-Q1 | **11,103** | *2,494* | *390* | *2,533* | *743* | **4,943** |
| 2006-Q2 | **11,054** | *2,474* | *387* | *2,427* | *782* | **4,984** |
| 2006-Q3 | **10,998** | *2,456* | *384* | *2,514* | *721* | **4,923** |
| 2006-Q4 | **10,814** | *2,389* | *377* | *1,873* | *777* | **5,398** |
| 2007-Q1 | **11,163** | *2,446* | *393* | *2,766* | *730* | **4,828** |
| 2007-Q2 | **11,078** | *2,425* | *387* | *2,673* | *691* | **4,902** |
| 2007-Q3 | **10,975** | *2,421* | *385* | *2,727* | *704* | **4,738** |
| 2007-Q4 | **10,791** | *2,380* | *379* | *2,147* | *698* | **5,187** |
| 2008-Q1 | **11,068** | *2,415* | *409* | *2,915* | *696* | **4,633** |
| 2008-Q2 | **11,002** | *2,407* | *404* | *2,804* | *681* | **4,706** |
| 2008-Q3 | **10,854** | *2,386* | *402* | *2,861* | *640* | **4,565** |
| 2008-Q4 | **10,650** | *2,358* | *392* | *2,208* | *741* | **4,951** |
| 2009-Q1 | **10,938** | *2,420* | *410* | *2,988* | *614* | **4,506** |
| 2009-Q2 | **10,837** | *2,411* | *409* | *2,839* | *572* | **4,606** |
| 2009-Q3 | **10,752** | *2,409* | *408* | *2,905* | *520* | **4,510** |
| 2009-Q4 | **10,601** | *2,435* | *401* | *2,179* | *613* | **4,973** |
| 2010-Q1 | **11,013** | *2,560* | *415* | *3,017* | *539* | **4,482** |
| 2010-Q2 | **10,982** | *2,593* | *410* | *2,880* | *556* | **4,543** |

## Appendix II – Variable Descriptions

### *Table A2- Variable Descriptions*

This table presents the definition of all the variables used in our models, including, when applicable, the formula applied to the original Compustat fields.

| Name | Definition | Compustat or Mnemonics |
|---|---|---|
| *Receivable Days* | Receivables / Daily Sales | rectrq / (saleq/ActDays) |
| *Payable Days* | Payables / Daily Cost of Goods Sold | apq / (cogsq/ActDays) |
| *Net Trade Credit (days)* | Receivable Days - Payable Days | recdaysA - paydaysA |
| *Receivables / Assets* | Receivables to Total Assets ratio | recdaysA = rectrq/atq |
| *Payables / Assets* | Payables to Total Assets ratio | paydaysA = apq/atq |
| *Sales / Assets (turnover)* | Sales to Total Assets ratio | salestoa = saleq/atq |
| *COGS to Assets* | Cost of Goods Sold to Total Assets ratio | cogstoa = cogsq/atq |
| *Inventory Days* | Inventory / Daily Cost of Goods Sold | invdaysA = invtq / (cogsq/ActDays) |
| *ActDays* | Actual number of days in each quarter | ActDays |
| *Gross Margin* | Gross Margin | smp = (saleq - cogsq)/saleq |
| *MP (pre)* | Market Power pre crisis (Lernex Index) proxied by gross margin average before crisis | smp |
| *HiMP (pre)* | Equals one if MP pre-crisis is above median and zero otherwise | hismp |
| *MP Census* | US Census Market Share of top 4 firms in 2007 for each NAICS 4 digit industry as in https://www.census.gov/econ/isp/ | mstop4 |
| *Hi MP Census* | Equals one if MP Census is above median and zero otherwise | himstop4 |
| *Cash / Assets* | Cash and Equivalents to Total Assets ratio | cashtoa = cheq / atq |
| *Short-term Debt / Assets* | Short-term debt to Total Assets ratio | stdtoa = dlcq/atq |
| *Long-Term Debt to Assets* | Long-term debt to Total Assets ratio | ltdtoa = dlttq/atq |
| *PPE / Assets* | Property, Plant and Equipment, net to Total Assets ratio | ppetoa = ppentq/atq |
| *Free Collateral* | PP&E after debt to total Assets ratio | freecoll = ( ppentq - dlttq) / atq |
| *Oper. Cash Flow / Assets* | EBITDA to Total Assets ratio | cftoa = oibdpq/atq |
| *Size* | natural log of Sales (millions) | lnsal = ln(saleq) |
| *Whited-Wu index* | Whited-Wu financial constraint index, Whited, T., & Wu, G. (2006). | ww = -0.091*cftoa - 0.062*dummydiv + 0.021*ltdtoa - 0.044* ln(atq) + 0.1021 * salesgrsic3 - 0.035*salesgr |
| *Hadlock-Pierce SA index* | Hadlock-Pierce financial constraint index, Hadlock, C., & Pierce, J. (2010). | sa = -0.737* ln(firmassetsMAX4_5) + 0.043*(ln(firmassetsMAX4_5)^2) - 0.040*ageMAX37 |
| *Industry Annual Sales Growth* | Annual SIC 3-digit industry growth | salesgrsic3 |
| *Firm Annual Sales Growth* | Firm annual growth | salesgr |

| | | |
|---|---|---|
| *Total Assets (top 4.5 billion)* | Total Assets limited to a maximum of $4.5 billion (as in Hadlock Pierce (2010)) | firmassetsMAX4_5 |
| *Firm's Age (top 37 years)* | Firm age since first time price appears in CRSP database, upper limit of 37 years. | ageMAX37 |
| *Dummy Dividend* | Equals one if firm paid dividend during the year (COMPUSTAT dvy) and zero otherwise | dummydiv |