Substitution Effects in Private Debt: Evidence from SMEs*

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Abstract

We investigate whether SMEs with demand for debt finance increase trade credit when they experience a negative shock to bank credit. We base our analysis on a large sample of SMEs from the five biggest EU countries. First, SMEs' ability to substitute largely depends on their credit quality. Second, substitution decreases during the financial crisis of 2007-09. Third, high credit quality firms with intermediate financial constraints are the most likely to substitute. We confirm these results on a subsample with matched bank-firm data. The evidence suggests that substitution in private debt is more difficult than considered in prior research.

JEL classification: G1, G20, G30, G32

Keywords: Bank loans, trade credit, asymmetric information, financial constraints, external finance dependence

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1. Introduction

Financing small- and medium-sized enterprises (SMEs) is challenging because these firms are more informationally opaque, risky, financially constrained, and bank-dependent than large firms. SMEs mainly rely on bank credit and trade credit from suppliers to raise external finance (e.g., Petersen and Rajan, 1994; Petersen and Rajan, 1997). They generally prefer bank credit over trade credit because the former tends to be less expensive and more flexible than the latter. But, how do SMEs respond when banks cut their lending? Is trade credit the appropriate response? Are there substitution effects between bank credit and trade credit? In this study, we investigate whether SMEs that have demand for debt finance increase trade credit (accounts payable) after a supply-side driven shock to their bank credit and which factors influence their response. We base the analysis on a large sample of SMEs from France, Germany, Italy, Spain and the U.K. and a subsample with matched bank-firm data.

The topic of our study is relevant for several reasons. First, SMEs represent a large fraction of all firms in many countries and contribute significantly to employment and growth. However, little is known about the dynamic interplay of bank credit and trade credit at individual SMEs. Second, variation in the availability of private debt can amplify or weaken the business cycle (e.g., King and Levine, 1993; Beck et al., 2000). If firms counter a negative shock to their bank credit by employing trade credit, they stabilize their access to credit throughout the cycle. However, if bank credit and trade credit are complementary (i.e., they increase or decrease at the same time), then booms and recessions are amplified, resulting in more volatile economic activity over time. Therefore, the question of whether the components of SMEs private debt finance are complements or substitutes are of first order importance. Third, if employing trade credit is not a sufficient response to a negative shock to bank credit supply, then policymakers should focus on stabilizing

bank credit supply and improving the bank lending environment in the first place, not consider trade credit as alternative to bank credit. SMEs can also stabilize bank credit supply at the micro level by combining forward and spot lending, different loan types and loans from institutions with different lending policies.

Researchers are examining the importance of trade credit provision (accounts receivable) and SMEs' use of trade credit (accounts payable) in different contexts, such as the effects of the recent financial crisis, trade credit chains, cost of capital, and economic growth (e.g., Garcia-Appendini and Montoriol-Garriga, 2013; Murfin and Njoroge, 2014; Jacobson and von Schedvin, 2015; Carbo-Valverde et al., 2016). The evidence is rather mixed. On the one hand, Carbo-Valverde et al. (2016) show that unconstrained firms in Spain mainly fund their investments with bank finance, while strongly constrained firms fund their investments with trade credit and this dependency increased during the recent financial crisis. Hence, trade credit can play a positive role because it helps to mitigate the adverse effects of a contraction of bank credit supply. On the other hand, Jacobson and von Schedvin (2015) analyze data from the Swedish credit bureau and highlight a negative role of trade credit. They document strong negative propagation effects in trade credit chains, arising from direct credit losses, negative credit quality spillover, and the loss of future business for suppliers.

Biais and Gollier (1997) and Petersen and Rajan (1997) suggest that there is a substitution relation between bank credit and trade credit. Substitution may occur because trade credit represents external finance for firms that are unable to attract sufficient bank credit because of severe informational asymmetries. The substitution hypothesis coincides with the redistribution view on trade credit (e.g., Love et al., 2007). Financially unconstrained firms redistribute part of

their bank credit to financially constrained client firms by providing trade credit (e.g., Garcia-Appendini and Montoriol-Garriga, 2013).

However, there are also arguments and evidence against a substitution relation between bank credit and trade credit. Bank credit creates a cash inflow that can be used for any purpose, while trade credit corresponds to a delayed payment for inputs for the production process. This fundamental difference might limit the possibility that these two forms of credit serve as substitutes for one another. Breza and Liberman (2015) find that trade credit does not create a special financing advantage but rather serves as mechanism to overcome contracting frictions between customers and suppliers. Murfin and Njoroge (2014) show the implicit costs of trade credit that have been neglected in earlier research. They focus on the retail industry and show that large, financially unconstrained buyers employ their bargaining power vis-à-vis small, financially constrained suppliers to obtain better terms in trade credit. Cuñat (2007) develops a model of trade credit and tests its predictions on a sample of firms from the U.K. He finds that trade credit insures firms against liquidity shocks and that it is mainly used when other forms of finance have been exhausted. These results are more pronounced when the links between supplier and customer are tight and the production process very specific. Fabbri and Menichini (2010) find that when the purchased goods are sufficiently liquid trade credit does not depend on the degree of financial constraints of the customer. The findings of these studies imply that trade credit is not a mere substitute for bank credit.

To identify a causal relation between a negative shock to SMEs' bank credit and their potential response in trade credit we apply a twofold strategy. First, we consider SMEs that have demand for external finance but experience a negative shock to bank credit supply. In this way, we study whether firms *can substitute* since we rule out the situation that firms *do not want to substitute*

because they have sufficient alternative funding sources available. Our setting allows us to interpret the negative complementary relation between bank credit and trade credit as evidence for firms' inability to substitute because we focus on external finance-dependent firms that have demand for external finance (Becker and Ivashina, 2014). Following the rationale of Rajan and Zingales (1998), we classify firm-year observations as external finance dependent if the cash flows are insufficient to cover the firm's investments. We measure the probability of substitution (firms exhibit an increase in trade credit after a negative shock to their bank credit) relative to a negative complementary relation (firms exhibit a decrease of trade credit and bank credit) with a novel firm-specific time-varying substitution indicator and study the factors that influence this probability. We base our analysis on a large dataset on SMEs from France, Germany, Italy, Spain, and the United Kingdom during 2006-2011. Second, we carry out an additional analysis with rich matched bank-firm data from Spain that allows us to establish a causal effect between the negative shock to SMEs' bank credit during the recent financial crisis and SMEs' ability to replace the funding gap with trade credit.

We obtain the following principal results. First, we find that substitution is more likely the higher the credit quality of the firm. Second, substitution decreased during the financial crisis of 2007-09 and further declined as the crisis deepened. Third, the impact of credit quality on the probability of substitution exhibits an inverse U-shaped relation with financial constraints. It is less important for the firms with low or no financial constraints because they are more likely to attract alternative forms of finance. It is also less important for the firms with high financial constraints because they are likely credit-rationed and therefore cannot borrow anyway. We confirm these results in a subsample analysis with matched bank-firm data from Spain.

Overall, the evidence suggests that trade credit is not the appropriate response to fill the funding

gap of SMEs when banks cut their lending. Firms with a lower credit quality have more difficulty to replace bank credit with trade credit, which contradicts the substitution hypothesis and the redistribution view. The probability of substitution decreased significantly when the recent financial crisis deepened. The substitution relation between bank credit and trade credit is time-varying and ultimately pro-cyclical, potentially amplifying the effect of recessions.

Our paper contributes to the SME finance literature in several ways. First, the substitution indicator makes it possible for us to examine the cross-sectional and time variation and allows us in combination with the other elements of our investigation (negative shock to bank credit and SMEs' external finance dependence), to identify a causal effect.

Second, we fill an important gap left by the study of Garcia-Appendini and Montoriol-Garriga (2013). We investigate SMEs' use of trade credit (accounts payable), whereas Garcia-Appendini and Montoriol-Garriga (2013) focus on relatively large U.S. firms from Compustat and investigate their provision of trade credit (accounts receivable). Because SMEs cannot access public debt markets, their external finance is largely limited to bank credit and trade credit. Thus, SMEs are the ideal testing ground to study substitution effects in private debt.

Third, we use data from the five biggest EU countries. Considering cross-country data is important because the firm characteristics, financial systems, and legal environments differ and affect the supply and demand for different types of debt (e.g., La Porta et al., 1997; Demirgüç-Kunt and Maksimovic, 2002; Beck et al., 2004; Berger and Udell, 2006; Haselmann et al., 2010).

The remainder of the paper is organized as follows. In Section 2, we present the related literature and develop our hypotheses. In Section 3, we describe the data and show summary statistics. In Section 4, we report the main results. In Section 5, we report additional analyses based on matched-bank firm data. We conclude in Section 6.

2. Related Literature and Hypotheses

2.1 RELATED LITERATURE

The literature has proposed several supply- and demand-side rationales as to why trade credit is a pervasive component on the firm's balance sheet. According to *supply-side* rationales, trade credit helps a firm to: (1) acquire private information about the customer (e.g., Smith, 1989; Mian and Smith, 1992; Petersen and Rajan, 1997); (2) enhance strong bargaining positions with customers (e.g., Giannetti et al., 2011; Cuñat, 2007); (3) decrease warehouse costs (Emery, 1987); and (4) facilitate a long-term supplier-customer relationship leading to future business opportunities (Ng et al., 1999).

The most important *demand-side* rationale for trade credit is that many firms, and in particular SMEs, resort to trade credit for financing because they are financially constrained and thus have limited or no access to other forms of external funding (e.g., Biais and Gollier, 1997; Petersen and Rajan, 1997). This rationale implies that bank credit and trade credit are substitutes for one another because these firms attract trade credit if they cannot secure bank credit. This interchangeability coincides with the redistribution view on trade credit (e.g., Love et al., 2007). The redistribution view implies that companies that borrow from financial intermediaries redistribute partly their borrowings to those who do not have access to financial intermediaries. The study of Garcia-Appendini and Montoriol-Garriga (2013) examines the provision of trade credit among large listed U.S. firms during 2005-2010 and confirms this mechanism.

Researchers further show that macroeconomic conditions have a significant influence on credit relationships (e.g., Petersen and Rajan, 1994; Berger and Udell, 2002; Nilsen, 2002; Ivashina and Scharfstein, 2010). Access to credit deteriorates during recessions because creditors become more

risk averse and therefore restrict credit extensions (Gertler and Gilchrist, 1994), implying that the relation between bank credit and trade credit varies with the state of the economy. A financial crisis will trigger liquidity shocks for certain firms, which will make them less creditworthy, thereby reducing their access to bank credit. This can lead to two opposite scenarios. The good scenario follows the substitution theory. In this case, firms experiencing credit constraints due to liquidity shocks will be able to secure more trade credit. The bad scenario holds that bank credit and trade credit exhibit a complementary relation. In this scenario, firms facing liquidity shocks will see their access to bank credit decline, along with their access to trade credit due to risk related to credit contagion (e.g., Jorion and Zhang, 2009). Deutsche Bundesbank (2012) and Kestens et al. (2012) document that trade credit extension decreased progressively during the recent financial crisis.

The common view in the earlier literature has been that trade credit (accounts payable) and short-term bank credit are substitutes for one another. However, there are arguments and evidence that challenge this view: (1) Firms cannot use trade credit in the same way as they can use bank credit. Trade credit does not create a cash inflow; it results from a delayed *cash outflow* for a specific purpose. In contrast, bank credit creates a *cash inflow* that can be used for any purpose. Hence, firms can use bank credit in a more flexible way than trade credit. (2) In many studies on trade credit, its interplay with short-term bank credit is not considered, nor is the potential endogeneity between trade credit and bank credit addressed. Uesugi and Yamashiro (2008) and Yang (2011) estimate the determinants of bank credit and trade credit in simultaneous equation models. The evidence provided by these studies casts doubt on the substitution hypothesis. (3) The common view that trade credit is less preferred than bank credit because of its higher price has been questioned in the recent literature. Giannetti et al. (2011) find that creditworthy firms with

some bargaining power in the market do obtain trade credit at a low cost. Fabbri and Menichini (2010) find that trade credit becomes cheaper than bank credit if the liquidation value of the purchased goods is relatively high. Relatedly, the redistribution view of trade credit implies that customers are weaker and more dependent than their suppliers. However, large customers usually have a significant bargaining power vis-a-vis smaller suppliers, indicating that suppliers can also be highly dependent on their customers (e.g., Giannetti et al., 2011; Albuquerque, Ramadorai and Watugala, 2015). (4) Many studies are based on single country data (e.g., Jacobson and von Schedvin, 2015; Boissay and Gropp, 2013; Kestens et al., 2012; Yang, 2011; Cuñat, 2007; Biais and Gollier, 1997; Petersen and Rajan, 1997). Haselmann et al. (2010) show that differences in the legal environment across countries influence credit markets. (5) Petersen and Rajan (1997) find that both the most and least profitable firms strongly use trade credit. (6) Fisman and Love (2003) find it is difficult for young firms to obtain trade credit. The substitution hypothesis implies that these firms cannot borrow from banks and therefore have to rely on trade credit.

2.2 HYPOTHESES

We derive a set of hypotheses on the determinants of the interplay between bank credit and trade credit (accounts payable). We note that the determinants of the probability of substitution between trade credit and short-term bank credit have not been directly examined in any of the related studies (e.g., Jacobson and von Schedvin, 2015; Carbo-Valverde, Rodriguez-Fernandez, and Udell, 2016; Garcia-Appendini and Montoriol-Garriga, 2013).

There are several firm characteristics that could affect the probability of substitution between bank credit and trade credit. We argue that credit quality is the most important factor to influence the probability of substitution. It should explain most of the cross-sectional and time variation in

individual firms' substitution behavior. We expect credit quality to have a positive impact on the probability of substitution because the higher the credit quality, the lower the default risk for the supplier and, as a result, the more likely that suppliers are willing to provide trade credit. This expectation is not trivial because trade credit is fully secured debt, implying that suppliers might be indifferent about the credit quality of their customer. In our analysis, we consider the reaction of firms in trade credit that experience a decrease in bank credit. We state Hypothesis 1 as follows:

Hypothesis 1: Credit quality has a positive impact on the probability of substitution between bank credit and trade credit.

Furthermore, we consider the time-series variation of the interplay between trade credit and bank credit. Earlier studies show that a financial crisis leads to a decrease in bank lending because banks are concerned about their liquidity and solvency (e.g., Ivashina and Scharfstein, 2010). Banks contract their credit supply in poor economic periods, resulting in an overall increase of financial constraints for firms. Highly constrained firms and firms with low credit quality might be the most affected by this mechanism. Yang (2011) shows that substituting bank credit for trade credit becomes more difficult during times of crisis due to credit contagion.

The impact that a financial crisis has on the interplay of bank credit and trade credit could vary during the different stages of the crisis. In the first stage of the recent financial crisis (i.e., the subprime mortgage crisis), banks had to build up capital that was eliminated by unprecedented losses. As a result, banks had to reduce lending, leading to a negative credit supply shock, as shown by Ivashina and Scharfstein (2010). After the collapse of Lehman Brothers (the second stage), the financial crisis evolved into a deep global recession, and bank lending problems got significantly worse. We state Hypothesis 2 as follows:

Hypothesis 2: The probability of substitution between bank credit and trade credit decreased

during the first stage of the recent financial crisis (H2a) and decreased more strongly during the second stage of the financial crisis (H2b).

Moreover, the impact of the firm's credit quality on the probability of substitution might interact with the firm's level of financial constraints. Firms that are either unconstrained or highly constrained are less likely to substitute sources of credit. We note that financial constraints and credit quality are related to each other, but not the same (e.g., Fazzari et al., 1988; Kaplan and Zingales, 1997). Firms that are financially constrained have difficulties in acquiring external finance because they are more opaque. Due to this opaqueness they have difficulties in signaling their true credit quality, which results in higher costs of debt. Therefore, the impact of the credit quality on the probability of substitution should be lower for unconstrained firms and for highly constrained firms. The first group is unlikely to substitute because they have access to alternative forms of external finance, while the second group is credit-rationed. The latter firms do not get any credit; lenders ignore differences in the risk of default of these firms. However, high credit quality firms with intermediate financial constraints might be able to substitute. Our hypothesis is partially supported by Atanasova (2007) who finds that financially constrained firms are more likely to resort to trade credit, while unconstrained firms avoid trade credit. A potential non-monotonic effect is not considered in that study. We state Hypothesis 3 as follows:

Hypothesis 3: Credit quality and financial constraints have a significant and non-monotonic interaction effect on the probability of substitution between bank credit and trade credit.

3. Data, Empirical Strategy and Variables

3.1 DATA SOURCE AND SELECTION CRITERIA

We collect firm data from the Orbis and SABI databases, both provided by Bureau van Dijk. These

databases contain firm-year observations from the five biggest countries in the EU (Germany, France, Italy, Spain, and the United Kingdom). Data for Spain comes from the SABI database, while the data from the other four countries are gathered from Orbis. We restrict our analysis to non-financial firms that are not publicly listed and that exhibit total assets not larger than €43 million in the last available year, consistent with the definition of SMEs from the European Commission (European Commission, 2005). Moreover, in Orbis there are many data points that report values of zero, potentially having an ambiguous meaning; they can either mean zero, "missing," or "unknown." To prevent this ambiguity in our dataset, we only include firms where the value of accounts payable, accounts receivable, and short-term bank credit equals at least €1,000 in any of the years in our sample period.

Applying these selection criteria results in our dataset with yearly data from 2006 to 2011 (2006 to 2010 for Spain). Since we use financial statement information we do not know the identity and number of firms' suppliers and banks. Nevertheless, we use a matched bank-firm dataset for Spain in an additional analysis.

In Orbis, the number of firms included in the database differs for each country, which results in certain countries being heavily over- or underrepresented in the raw dataset. Therefore, we construct the dataset in a way that gives each country a weight that is proportional to its average GDP over the sample period. The final dataset is comprised of 1,186 SMEs from Germany (28%), 922 from France (22%), 920 from the U.K. (21%), 751 from Italy (17%), and 501 from Spain (12%). In that sample, we include the largest SMEs from each country. To rule out that our results are driven by a selection bias, we also stratify the raw sample in size quintiles and randomly draw a number of firms (equal to the number of firms for each country in the main dataset divided by five) within each country-quintile and repeat this procedure 100 times.

3.2 EMPIRICAL STRATEGY

A major challenge in studying the relation between bank credit and trade credit at the firm level is that both variables may change simultaneously. The resulting potential endogeneity has not been addressed in many studies (e.g., Kestens et al., 2012; Nilsen, 2002) because they either regress trade credit on bank credit (or the other way around) or do not control for bank credit, making it difficult to draw conclusions about the complementary or substitution relation between trade credit and bank credit.

We focus on the probability of substitution between the components of SMEs' private debt and its determinants, while earlier studies examine the determinants of trade credit or bank credit (e.g., Garcia-Appendini and Montoriol-Garriga, 2013; Petersen and Rajan, 1997). The substitution indicator (SI_{it}) equals zero for a negative complementary relation between changes in bank credit (ΔB_{it-1}) and changes in trade credit (ΔT_{it}), and one for a substitution relation between changes in bank credit and changes in trade credit, both conditional on a negative shock to short-term bank credit in the previous year. It is defined as shown in Equation (1)[1]. In this specification we define trade credit as the accounts payable.

[1] In a robustness test we consider a modified substitution indicator that takes also contemporaneous shocks to bank credit into account (Section 4.5 and Table IX).

$$SI_{it} = \begin{cases} 0 \text{ if } \Delta B_{it\text{-}1} < 0 \cap \Delta T_{it} < 0 \\ \\ 1 \text{ if } \Delta B_{it\text{-}1} < 0 \cap \Delta T_{it} > 0 \end{cases} \quad (1)$$

Our identification strategy is twofold. First, we focus on firms that have demand for external finance to ensure that they want to substitute bank credit for trade credit. To identify firms with demand for credit, we use the concept of external finance dependence proposed by Rajan and Zingales (1998). Firms that have no demand for external finance do not need to substitute bank credit because they have sufficient internal finance to fund their operations (Becker and Ivashina, 2014; Duchin et al., 2010). External finance dependence is calculated as formulated in Equation (2):

$$EFD_{it} = \frac{\Delta TA_{it} - CF_{it}}{\Delta TA_{it}} \quad (2)$$

ΔTA_{it} is a proxy for a firm's yearly level of investments and CF_{it} represents the firm's annual cash flows. Only observations with a positive outcome are included in the analyses because these are the firms that theoretically need credit.[2] We calculate EFD_{it} at the firm level and, alternatively, at the industry-country level. For the latter, we use the median values of ΔTA_{it} and CF_{it} at the industry-country level. On the one hand, EFD_{it} at the firm level is more informative about a firm's specific needs for external finance than EFD_{it} at the country-industry level. On the other hand, the level of

[2] Alternatively, we computed *EFD* only for observations with $\Delta TA > 0$ because *EFD* becomes positive for $\Delta TA < 0$ and $CF > 0$. The results are similar.

investments depends on a firm's access to finance, making EFD_{it} potentially endogenous at the firm level, but not at the country-industry level. Because of these reasons we decided to consider both measures of $EFD_{it.}$

We note that our sample period of 2006-2011 includes a severe macroeconomic recession, during which many banks had to reduce their lending significantly because of large losses, illiquidity, and insolvency concerns (Ivashina and Scharfstein, 2010; Duchin et al., 2010; Puri et al., 2011). Therefore, we study the response of trade credit after a negative shock to SMEs' bank credit, which is for the vast majority of SMEs an exogenous and credit supply-side driven shock.

Second, we conduct an additional analysis with rich matched bank-firm on Spain that allows a direct identification of firms that were facing a negative shock to bank credit during our sample period. We distinguish between SMEs that have a relationship with an unhealthy bank (measured by the incidence of being bailed out by the government) and firms that do not have a relationship with an unhealthy bank. SMEs that borrow from unhealthy banks were facing a stronger reduction in credit supply and therefore had a higher need for substitution during the recent financial crisis. We examine whether trade credit helped these firms to replace the funding gap.

3.3 VARIABLES AND SUMMARY STATISTICS

Our dependent variable is the substitution indicator SI_{it}, as explained above. Figure 1 shows the relative frequency of substitution (SI_{it}=1) by country and over time. There is a sharp decrease in the fractions of substitution relations throughout the financial crisis until 2009, as well as an increase during the recovery in 2010. All five countries show a similar pattern but the effects vary in terms of their magnitude. The mean of SI in the entire sample equals to 0.49, indicating that substitution and complementary relationships are on average almost equally likely to occur.

However, the overall mean value clouds that there is substantial variation over time (e.g., the yearly mean of SI in the U.K. changes from 0.28 in 2008 to 0.65 in 2010) that indicates that the probability of substitution depends on the state of the economy.

(Insert Figure 1 here)

The explanatory variables are the potential factors that might influence the substitution indicator as posited in Hypotheses 1-3. The main explanatory variable is the credit quality of the firm, which we measure with the Altman Z-score (Z) for private firms (Altman, 1968). The Z-score is a widely used composite measure of credit quality (firm default risk) and is based on several factors, such as liquidity, retained earnings, profitability, leverage, sales, and size. Agarwal and Taffler (2007) show that the Z-score is predictive of the default risk of firms in different time periods and different countries. Altman's Z-score[3] for private companies is computed as shown in Equation (3). All components are winsorized at the 1st and 99th percentile to ensure that the Z-score is not driven by extreme observations.

$$Z_{it} = 0.7\frac{WorkingCapital_{it}}{TotalAssets_{it}} + 0.85\frac{RetainedEarnings_{it}}{TotalAssets_{it}} + 3.1\frac{EBIT_{it}}{TotalAssets_{it}} + 0.4\frac{TotalAssets_{it}}{TotalLiabilities_{it}} + \frac{Sales_{it}}{TotalAssets_{it}}. \quad (3)$$

As stated in Hypothesis 3, the influence of Altman's Z-score on the probability of substitution might be dependent on the level of financial constraints of the firm. We measure financial constraints with a widely used measure, the KZ index (Kaplan and Zingales, 1997; Lamont, Polk,

[3] Sales are not available for firms from the U.K. We therefore use operating revenues in all countries. For EBIT, we take ROA before taxes instead. Retained earnings are not directly available in Orbis. We have estimated them as equity minus capital (firm wealth minus the value of the shares).

Saá-Requejo, 2001), which is defined in Equation (4).[4] All components are winsorized at the 1st and 99th percentile. In order to measure the non-monotonicity between the Z-score and the KZ index, we create quintile dummies for the latter (*KZ_Q*).

$$KZ_{it} = -1.002\frac{CF_{it}}{TA_{it-1}} + 3.139\frac{TL_{it}}{TA_{it-1}} + 39.368\frac{Div_{it}}{TA_{it-1}} - 1.315\frac{Cash_{it}}{TA_{it-1}}. \quad (4)$$

Because there has been debate about how to measure financial constraints (e.g., Farre-Mensa and Ljungqvist, 2016), we consider the WW index (Whited and Wu, 2006) and the SA index (Hadlock and Pierce, 2010) as alternative measures in robustness tests.

Moreover, we examine the impact of the recent financial crisis, using a set of dummy variables that indicate different stages of the crisis. In continental Europe, the first (second) stage of the crisis, *D_Crisis1* (*D_Crisis2*), is a dummy variable equal to one in the year 2008 (2009), and zero otherwise. We consider 2009 as the second stage of the crisis because Lehman Brothers collapsed in September 2008, which is the trigger event of the deep global recession. In the U.K., we consider 2007 (2008) as the first (second) stage of the crisis because the crisis started earlier and evolved faster in the U.K. due to its stronger ties with the U.S. The post-crisis period is indicated by a dummy variable (*D_Aftermath*) that equals one in the years 2010-2011 (2009-2011 for the U.K.) and zero otherwise.

We add several control variables that might influence the substitution indicator. The first variable is firm size (*LnTA*), measured by the natural logarithm of total assets. The second variable is collateral, motived by the study of Cuñat (2007), and measured in two ways. Long-term

[4] Dividends are not available in Orbis. We estimate dividends as net income minus the change in equity (i.e., the proportion of income that is not retained by the company).

collateral is measured with fixed tangible assets (*TangFA*) and short-term collateral with inventories (*Inv*) (e.g., Norden and van Kampen, 2013), both scaled by total assets. The third variable is the sum of cash and cash equivalents divided by total assets (*Cash*). The last variable is profitability measured by ROA. In all regressions, we control for industry and country fixed effects, where industry is derived from the two digit SIC code. Industry fixed effects are important because suppliers are more willing two extend trade credit to customers in industries with high product specificity (Cuñat, 2007). Country fixed effects are important because heterogeneity in financial and legal systems creates heterogeneity in financial markets (e.g., La Porta et al., 1997; Haselmann et al., 2010). Z and ROA are highly sensitive to outliers and are winsorized at the 1st and 99th percentile at the country level.

(Insert Table I here)

The summary statistics in Table I are reported separately for the firms that exhibit a complementary relation (column (0)) and those that exhibit a substitution relation (column (1)). Panel A reports the descriptive statistics for *all firms* experiencing a negative shock in bank credit, while Panel B reports the statistics for the *firms that have demand for external finance* (EFD firms). The mean and median values of the Z-score are higher for the substitution firms, indicating that firms of higher credit quality substitute credit more often. The mean and median of the Z-score differ substantially across countries (not reported); German firms have the highest credit quality with a mean (median) Z-score of 3.36 (3.13), while Italian firms have the lowest credit quality with a mean (median) Z-score of 1.78 (1.68). The other three countries have Z-scores between 2 and 3. In addition, the value for Z (KZ) drops (rises) when we exclude firms that have no demand

for external finance. This is intuitive because firms that do not need external finance display usually a higher credit quality and lower financial constraints.

4. Empirical Analysis

4.1 BASELINE ANALYSIS

First, we investigate which factors influence whether SMEs increase trade credit after they have experienced a negative shock to their bank credit by regressing the SI_{it} on the lagged Z-score, the crisis dummies, and lags of the control variables. Table II presents our baseline results.

(Insert Table II here)

We find that the credit quality measure $Z_{it\text{-}1}$ has a consistently positive impact on the probability of substitution. As shown in Table II, a one-unit increase in $Z_{it\text{-}1}$ is associated with a 7.8% increase in the probability of substitution. The result indicates that credit substitution for low credit quality SMEs is difficult because suppliers care about the ex ante default risk of their customers although trade credit is fully secured debt. Moreover, the probability of substitution went down significantly during the crisis: it is 28.1% lower during the first stage of the financial crisis and 52.0% during the second stage compared to the pre-crisis level, respectively. These findings remain similar if we add control variables in column (2) and become even stronger if we only consider firms that have demand for external finance in column (3). We note that the sample size decreases only slightly when we exclude firms that do not depend on external finance (from 8,825 to 7,040). This is not surprising because SMEs are in general strongly dependent on external finance. In column (3), a one-unit increase in $Z_{it\text{-}1}$ is associated with a 9.9% increase in the probability of substitution. We

thus conclude that, all else being equal, the relation between credit quality and substitution is monotonic. In column (4), we present the results of a regression when the variables are demeaned at the country-level median and obtain similar results. In column (5) we report the results for firms located in external finance-dependent industries. A one unit increase in Altman's Z increases the probability of substitution with 6.7%, the result becomes weakly significant though (p=0.068). The impact of the crisis dummies stay similar in both economic and statistical significance. Overall, the results are consistent with Hypotheses 1 and 2.

We repeat the baseline analysis on a country-by-country basis for the external finance-dependent firms to ensure that our effects are not driven by one particular country. The results are presented in Table III. For this analysis, we employ the larger raw samples from each country to make full use of the data (we do not use the aggregate sample based on GDP shares, as explained at the end of Section 3.1).

(Insert Table III here)

The findings in Table III confirm our aggregate analysis shown in Table II and are consistent across all five countries. However, the magnitude of the effects varies substantially. The Z-score has a positive and significant impact on the probability of substitution. A one unit increase in the Z-score is related to an increase in the probability of credit substitution between 8.2% (Germany) and 18.4% (France). The effect is the smallest in Germany because its financial system was less strongly hit by the financial crisis and long-term bank credit to SMEs is more common than in the other four countries. Also the results for the financial crisis are consistent for all countries, confirming that substitution is strongly dependent on the state of the economy as shown in Figure

1. The probability of credit substitution decreased significantly in both stages of the financial crisis in all five countries; the second stage of the crisis always has an even more negative impact than the first stage of the crisis. The probability of credit substitution during the second stage of the crisis decreases between 36.5% (Italy) and 72.6% (U.K.). In additional analyses[5] we find that the results are similar for the Altman's Z-score and D_Crisis2 when we investigate firms from external finance-dependent industries: D_Crisis2 is significant in all five countries and the Altman's Z-score in four countries (Germany is the only exception). Furthermore, we estimated the baseline model with interaction effects between the explanatory variables and country dummies on the (unweighted) aggregate sample that consists of the country-specific raw samples. The results are in line with the ones we report in Table III. We conclude that the effects of credit quality and the crisis on the substitution of bank credit for trade credit after a negative shock to bank credit are qualitatively similar but the magnitudes vary across countries.

In the next step, we investigate how much SMEs could substitute, using a modified version of the substitution indicator: the three-outcome $SI3_{it}$.[6] This modified version has three possible outcomes: (1) negative complementary relation; (2) partial substitution; and (3) perfect substitution. Partial substitution refers to the situation where trade credit increases in year t to a lower extent than bank credit decreased in year t-1, while perfect substitution refers to the situation where trade credit increases in year t at least as much as bank credit decreased in year t-1. In other words, we investigate two different forms of substitution; the situation where firms do not fully fill the funding gap resulting from the decrease in bank credit and the situation where they do fully fill this gap. Table IV presents the results. We estimate the probability of partial or perfect credit

[5] Results are available from the authors on request.

[6] We also considered the elasticity of trade credit to bank credit as an alternative version of the substitution indicator. However, it turned out that the elasticity measured at the firm level is too volatile. The discrete SI3 indicator is more robust and implicitly depends on the elasticity as input.

substitution relative to the probability of a negative complementary relation.

(Insert Table IV here)

We find in Table IV that perfect credit substitution is more likely the higher the credit quality of the firm. For firms with external finance dependence at the firm (industry) level a one-unit increase in the Altman Z-Score increases the probability of substitution with 17.4% (15.1%). Furthermore, substitution decreased in both the first and second stages of the financial crisis. There is a sharp drop in substitution in the second stage of the crisis (odds ratios of 0.35-0.55 in most specifications). By and large, these results are consistent with those from Table II and III.

4.2 CREDIT QUALITY AND STAGES OF THE CRISIS

We have shown that the probability of substitution went down significantly during both stages of the financial crisis, but it declined more strongly during the second stage, as posited in Hypothesis 2. We now investigate in Table V whether the impact of credit quality on the probability of substitution varies during the different stages of the financial crisis. We perform this analysis for the full sample and the sample of firms/industries that have demand for credit (EFD firms/industries).

(Insert Table V here)

We obtain three results. First, there is a significant and negative impact of the second stage of the crisis on the probability of substitution in all samples. Hence, after the failure of Lehman

Brothers in fall 2008 the probability of substitution decreased significantly compared to the pre-crisis period. Second, for the full sample the interaction term between Z_{it-1} and *D_Crisis1* is significantly negative, suggesting that high credit quality firms are less likely to substitute bank credit for trade credit. This finding is plausible for the full sample where we do not limit the analysis to firms that have demand for credit. High credit quality SMEs are likely to have alternative sources of debt available (e.g., credit from the owners, friends or family) during the first stage of the financial crisis, likely to retain sufficient earnings to fund their operations, or they simply shrink by decreasing leverage. Consistent with this reasoning, this effect disappears in the analysis for EFD firms. Third, for the sample of firms with demand for credit (EFD firms, both at the firm and industry level) the interaction term between Z_{it-1} and *D_Crisis2* has a significantly negative impact on the probability of substitution, suggesting that even high credit quality firms had difficulties in compensating for the negative shock to bank credit supply with trade credit. Hence, suppliers' willingness to lend shows a similar pattern as the one of banks. Moreover, the findings are consistent with the fact that trade credit insurers reduced their coverage, increased retention amounts and risk premiums when the crisis deepened (e.g., Deutsche Bundesbank, 2012).

4.3 CREDIT QUALITY AND FINANCIAL CONSTRAINTS

It is possible that the observed effect of credit quality on the probability of substitution is non-monotonically related to the level of financial constraints, as stated in Hypothesis 3. The concept of financial constaints is related but not the same as financial distress, as pointed out in the literature (e.g., Fazzari, Hubbard and Petersen, 1988; Kaplan and Zingales, 1997). To analyze this issue, we interact the Z-score (Z_{it-1}) with the KZ index quintile dummies (KZ_Q_{it}), using the first quintile as reference category. Table VI reports the results.

(Insert Table VI here)

The regression results indicate an inversely U-shaped relation between the probability of credit substitution and the interaction term for the full sample (column 1) and the sample of the external finance dependent firms (column 2). The finding indicates that credit quality is most important for the firms with intermediate financial constraints. For firms in quintile 3 (4), a one-unit increase in the Z-score of EFD firms increases the probability of substitution by 15.1% (23.8%) relative to quintile 1. The effect is statistically significant but less pronounced for the firms with no financial constraints because these firms have access to alternative forms of finance. In contrast, the Z-score matters less for the firms with the highest financial constraints because these firms are likely subject to credit rationing.

In addition, we repeated the regression with the quintile dummies for the WW index. The results for the interaction effects with the WW index are similar to those for the KZ index, indicating that the probability of substitution exhibits the highest sensitivity to credit quality for the firms in the fourth quintile. We also considered the SA index (Hadlock and Pierce, 2010), but we do not find any significant interaction effects of this index with credit quality. One explanation is that the SA index does not sufficiently discriminate between the firms in our sample because they are all relatively small.

Finally, we conduct the baseline regression from Table II for external finance-dependent firms separately for each KZ quintile by country. The odds ratios of the Z-score in each KZ quintile group in the raw sample are plotted in Figure 2. All five EU countries display an inversely U-shaped pattern. The results are consistent with Hypothesis 3, confirming that firms with a high

credit quality and intermediate financial constraints exhibit a higher probability of credit substitution than others.

(Insert Figure 2 here)

4.4 SUBSTITUTION BETWEEN TOTAL BANK CREDIT AND TRADE CREDIT

In the previous analyses, we investigated SMEs' reaction with trade credit in year *t* after a negative shock to their short-term bank credit in year *t*-1. On the one hand, it is possible that SMEs were also facing a negative shock to their long-term bank credit, especially those firms that had long-term bank credit expiring during the recent financial crisis (Campello et al., 2012). On the other hand, it is most likely that firms substitute short-term bank credit (and not long-term bank credit) with trade credit (and vice versa). In other words, it is unlikely that firms substitute a permanent drop in long-term bank credit with a permanent increase in trade credit as the purpose and the cash flow effects of these two types of debt finance are very different.

To provide further evidence on the substitution effects between bank credit and trade credit, we examine the response of trade credit in year *t* after a negative shock to total bank credit (short-term and long-term bank credit) in year *t*-1 (SI_{it}^{total}).

(Insert Table VII here)

When comparing the results for the shock to total bank credit reported in Table VII with the baseline results from Table II, we see that the Z-score and the financial crisis have a similar impact on the probability of credit substitution. For external finance-dependent firms, a one-unit increase

in the Z-score increases the probability of substituting total (short-term) bank credit for trade credit by 7.7% (9.9%). In addition, the probability of substituting total (short-term) bank credit decreases 34.0% (32.9%) during the first stage of the financial crisis and 55.2% (55.9%) during the second stage. The fact that the results do not change when we use long-term debt as input for the substitution indicator increases the reliability of the evidence because trade credit is a better substitute for short-term bank credit than for long-term bank credit because of its maturity match.

4.5 MODIFIED VERSIONS OF THE SUBSTITUTION INDICATOR

In the above analyses we investigated SMEs' substitution conditional on a *negative* shock to bank credit. We now consider two modified versions of the substitution indicator. First, to rule out that our results are driven by a selection bias, we repeat the baseline analysis with a substitution indicator that is unconditional on the nature of the shock to bank credit in year t-1 ($SI4_{it}$) as shown below. In other words, we now consider *positive and negative shocks* to SMEs' bank credit in year t-1 to study their response in trade credit.

$$SI4_{it} = \begin{cases} 1 \quad if\ \Delta B_{it-1} < 0 \cap \Delta T_{it} < 0\ (complementary) \\ 2 \quad if\ \Delta B_{it-1} < 0 \cap \Delta T_{it} \geq 0\ (substitution) \\ 3 \quad if\ \Delta B_{it-1} \geq 0 \cap \Delta T_{it} < 0\ (substitution) \\ 4 \quad if\ \Delta B_{it-1} \geq 0 \cap \Delta T_{it} \geq 0\ (complementary) \end{cases} \tag{6}$$

The results of the multinomial regression are presented in Table VIII. All probabilities are relative to having a negative complementary relation. In the least restrictive sample, we find that a one-unit increase in Altman's Z increases (decreases) the probability of substituting bank credit (trade credit) for trade credit (bank credit) with 8.4% (10.3%). This means that firms with higher credit quality are more likely to use trade credit when bank credit is unavailable and that these firms are less dependent on trade credit when bank credit is available. Furthermore, we show that the occurrence of a crisis decreases the probability of being in group 2, 3 or 4. In other words, a negative complementary relation is most likely to occur during a crisis.

(Insert Table VIII here)

Second, the substitution indicator in Equation (1) does not take into account a potential change in bank credit in year *t*. However, this could impact the probability of substitution because if a decrease in bank credit in the previous year is fully offset by an increase in bank credit in the current year, then substitution is not necessary (and hence a complementary relation becomes more likely). For this reason, we additionally condition the substitution indicator on either $\Delta B_{it} < 0$ or $|\Delta B_{it}| < |\Delta B_{it-1}| \cap \Delta B_{it} > 0$. In other words, we only include cases where the negative shock to bank

credit in year t-1 is not fully offset by a positive shock in year t. The results are presented in Table IX and qualitatively confirm our previous findings.

(Insert Table IX here)

4.6 STRATIFIED RANDOM SAMPLING

The previous analyses are based on a sample comprised of the largest SMEs from the five biggest EU countries. To rule out that there are selection effects due to firm size, we stratified our raw sample using firm size quintiles. For each country we select a number of firms by random sampling within each size quintile. This number equals the number of firms per country in the raw sample divided by five. We repeat this procedure 100 times. We then re-estimate our main regressions for firms that are externally financially dependent for each of the 100 subsamples and report the mean regression statistics in Table X.

(Insert Table X here)

For the regression corresponding to Table II, we find in Table X that the Z-score has a significantly positive impact on the probability of credit substitution, where the mean odds ratio is 1.113 (p=.000). In addition, in all 100 reiterations, the first and second stages of the financial crisis have a significantly negative impact on the probability of credit substitution, where the magnitude of the second stage is always larger than the first stage. The mean odds ratios are 0.681 (p=.000) and 0.478 (p=.000), respectively. When we repeat the analysis for the regression corresponding to Table V, we obtain similar results on both the Z-score and the crisis dummies. The mean odds

ratio of the Z-score is 1.123 (p=.000) and is positively significant in all 100 subsamples. For the first and second stages, the mean odds ratios are 0.685 (p=.024) and 0.549 (p=.000), respectively. The crisis dummies have a significant negative effect in 79 and 98 of the subsamples, respectively. The interaction terms between the Z-score and both crisis dummies are insignificant though. In the regression corresponding to Table VI and Figure 2 (not reported here to conserve space), we confirm the inversely U-shaped pattern for the interactions between the KZ index quintiles and the Z-score after the stratified random sampling.

Overall, our baseline results from Table II, V and VI remain robust when we repeat the analysis on the stratified random sample.

5. Analysis with Matched Bank-Firm Data

We carry out one more analysis that complements our previous evidence. The following analysis allows us to directly identify a causal relation between a decrease of bank credit supply in year t-1 and the potential response in trade credit in year t at the individual firm level. We do so by using rich matched bank-firm data on SMEs from Spain.[7]

For each firm and year, we observe the number and names of banks from which the SMEs obtain credit, resulting in a dataset comprised of 59,534 bank-firm-year observations for the period 2005-2010. The changes in credit we observe in this dataset are effective changes in credit used by firms. This approach is consistent with the earlier analysis based on the substitution indicator that captures firms' response to a negative shock to bank credit supply.[8]

[7] For the other EU countries we cannot use matched bank-firm data because either the information on the firms' bank relationships is not available in ORBIS or it is time-invariant and therefore not reliable.

[8] It would be interesting to distinguish binding and non-binding shocks to bank credit supply. Unfortunately, information on unused lines of credit is not available in our database. We focus on binding shocks to bank credit supply, as the substitution indicator is based on changes in bank credit and trade credit.

To ensure that firms did not decrease their borrowings voluntarily (credit demand-side effect), but that banks cut lending to the firms (credit supply-side effect), we collect information on bank bailouts in the Spanish banking sector. The indicator variable *Bailout* equals one if the bank was eventually bailed out during the recent financial crisis, and zero otherwise. We follow the Bank of Spain and consider all types of government intervention: full bailout, capital infusion, debt guarantees, and other instruments.[9] The Spanish banks that were eventually bailed out, which are almost all the savings banks and some commercial banks, started to experience dramatic losses in 2007-2008. This situation arose due to the collapse of the Spanish housing bubble and the resulting losses from domestic mortgage lending and banks' credit exposure to securitized U.S. subprime mortgages, which forced these banks to significantly reduce their lending activities (Illueca et al., 2014). This decrease in lending of bailed out banks was significantly stronger than that of banks that were not bailed out. Firms borrowing from bailed out banks faced an exogenous supply-side driven shock to their bank credit. Our identification strategy follows Puri et al. (2011), who study the change in rejection rates of German savings banks that are connected with Landesbanks that were (or were not) affected by the U.S. subprime mortgage crisis. We consider only firms having demand for credit, i.e., those that are considered as external finance dependent (Rajan and Zingales, 1998).

The *Bailout* dummy is an ex-post indicator for banks that were forced to cut lending due to financial distress. However, being bailed out by the government is a result of fundamental financial problems. Therefore, it is reasonable to assume that the bailout banks had to cut lending during this time period. We also create a dummy indicator for savings banks that serves as an ex ante indicator of unhealthy banks. Savings banks in Spain were taking substantial risks before the 2008

[9] Details on the total amount of public money injected into troubled banks are available at the Bank of Spain's website: http://www.bde.es/f/webbde/GAP/Secciones/SalaPrensa/NotasInformativas/Briefing_notes/es/notabe040515.pdf.

financial crisis through aggressive loan growth. Their loan growth came to an abrupt end in 2007-2008. As consequence of substantial losses, the savings banks were forced to cut lending, resulting in a significant credit crunch for SMEs in Spain.

We estimate the probability of credit substitution with the indicator variables for the financial crisis (*Crisis*), for either bailed out banks (*Bailout*) or savings banks (*Savingsbank)*, the interaction term (*Crisis*Bailout*) or *(Crisis*Savingsbank)*, firm controls, bank controls, and industry fixed effects. We cluster the standard errors at the bank-firm level. Table XI reports the odds ratios and *p*-values from three specifications of the logit regression model.[10]

(Insert Table XI here)

This analysis yields a clear result. The odds ratios of the interaction term *Bailout*Crisis* is significantly below one in all three specifications in Table XI. This finding indicates that firms borrowing from bailed out banks during the recent financial crisis had a lower probability of credit substitution than other firms. The effect is significant at the 1% level for the baseline model reported in column (1); it remains significant at the 1% level when we add a comprehensive set of time-varying firm and bank controls in columns (2) and (3) and when we change the *Bailout* dummy into the *Savingsbank* dummy in column (4). We also find that the probability of credit substitution significantly decreases during the financial crisis, as indicated by odds ratios for the variable crisis that range between 0.71 and 0.88. Moreover, the odds ratio for the firms' Z-score is above one and significant, confirming our earlier result that higher credit quality increases the

[10] An alternative model specification, in which the bailout indicator is defined as the ratio of total amount of funds injected by the government to total banks' equity at the beginning of the financial crisis, leads to similar results. This alternative analysis is available from the authors upon request.

probability of substitution. These results are consistent with the previous ones for the five largest EU countries and the aggregate sample, as shown in Table II and III.

We consider the analysis with matched bank-firm data as an additional piece of evidence for a causal effect: SMEs with demand for credit found it difficult during the recent financial crisis to (sufficiently) replace the drop in bank credit with trade credit. In other words, trade credit did not help to fill SMEs' funding gap caused by the negative shock to bank credit supply.

6. Conclusion

We investigate whether SMEs that have demand for debt finance use trade credit when they experience a negative shock to their bank credit and which factors influence their response over time and across countries. We base our analysis on a large sample of SMEs from France, Germany, Italy, Spain and the U.K.

We find that substitution and complementary relationships between bank credit and trade credit are on average equally likely during 2006-2011, but their importance varies substantially over time and across countries. Firms with higher credit quality are generally more likely to substitute credit. Substitution became less likely during the financial crisis and it further declined as the crisis deepened. High credit quality firms with intermediate financial constraints are the ones that are most likely to substitute. We carry out an additional analysis with matched bank-firm data and find that the SMEs that are hit more by the shock to bank credit supply are less likely to substitute.

We show that the probability of credit substitution depends on firms' credit quality, financial constraints, macroeconomic conditions, and country effects. Our study provides evidence on the dynamic interplay of the components of private debt in SME finance and has implications for the institutional design of the lending environment, economic policy and individual firm behavior.

The evidence suggests that substitution between the components of private debt is more difficult than considered in prior research. Trade credit is *not* the appropriate response to fill the funding gap that emerges when banks cut lending to SMEs. Policymakers should focus on enhancing financial stability, and thereby stabilize bank credit supply and the bank lending environment, rather than considering trade credit as alternative mode of external finance to mitigate the adverse effects on the real economy. SMEs can stabilize their access to bank credit by combining forward lending and spot lending and by diversifying across loan types and financial institutions.

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Table I. Summary statistics

This table reports descriptive statistics for all non-indicator variables. We present the number of firm years for each variable and the mean, median, and standard deviation for both outcomes of SI_{it}. Panel A presents the descriptive statistics for all firms that were facing a negative shock to bank credit in year t-1, while Panel B presents the descriptive statistics only for the firms that are externally finance-dependent (hereinafter: EFD firms; the increase in total assets exceeds the value of operating cash flows, as is proposed by Rajan and Zingales 1998).

Panel A: All firms

	Firm-Years	Mean		Median		St. Dev.	
		(0)	(1)	(0)	(1)	(0)	(1)
Z	9,222	2.72	2.89	2.58	2.71	1.36	1.47
KZ	8,808	-0.24	-0.12	0.01	0.29	3.45	3.33
KZ_Q	8,808	2.85	3.03	3.00	3.00	1.40	1.42
Size	9,672	9.63	9.44	10.14	9.98	1.22	1.25
Cash	9,434	0.08	0.08	0.03	0.03	0.12	0.11
Inventories	9,570	0.20	0.20	0.14	0.14	0.22	0.21
Tangibles	9,618	0.27	0.28	0.20	0.20	0.26	0.26
ROA	9,392	0.03	0.03	0.02	0.02	0.08	0.08

Panel B: External finance-dependent firms (EFD firms)

	Firm-Years	Mean		Median		St. Dev.	
		(0)	(1)	(0)	(1)	(0)	(1)
Z	7,365	2.65	2.86	2.49	2.66	1.35	1.50
KZ	7,011	-0.22	0.09	0.06	0.45	3.49	3.33
KZ_Q	7,011	2.88	3.13	3.00	3.00	1.41	1.42
Size	7,746	9.68	9.42	10.19	9.98	1.24	1.29
Cash	7,561	0.07	0.07	0.03	0.03	0.11	0.11
Inventories	7,663	0.21	0.22	0.15	0.15	0.22	0.22
Tangibles	7,693	0.27	0.26	0.19	0.18	0.26	0.26
ROA	7,479	0.02	0.03	0.02	0.02	0.07	0.07

Table II. The determinants of the probability of substitution

This table reports results from the logit regression analyses where SI_{it} is regressed on the Z-score, the different stages of the crisis, cash holdings, tangible fixed assets, inventories, firm size, ROA and a set of country and industry dummies. This regression analysis informs how the explanatory variables increase or decrease the probability of substitution between short term bank credit and accounts payable. We report the odds ratios with the *p*-values in parentheses for each explanatory variable. ***, **, * indicate coefficients that are statistically significant at the 1%, 5%, and 10% levels, respectively, using robust standard errors clustered at the firm level.

	(1)	(2)	(3)	(4)	(5)
	Full Sample	Full Sample	EFD Firms	EFD Firms demeaned	EFD Ind.
Z(t-1)	1.078	1.087	1.099	1.093	1.067
	(0.000) ***	(0.000) ***	(0.000) ***	(0.000) ***	(0.068) *
D_Crisis1	0.719	0.724	0.671	0.634	0.581
	(0.000) ***	(0.000) ***	(0.000) ***	(0.000) ***	(0.000) ***
D_Crisis2	0.480	0.485	0.441	0.410	0.492
	(0.000) ***	(0.000) ***	(0.000) ***	(0.000) ***	(0.000) ***
D_Aftermath	1.084	1.084	1.014	0.933	1.050
	(0.222)	(0.232)	(0.852)	(0.345)	(0.657)
Size(t-1)		0.900	0.876	0.882	0.628
		(0.000) ***	(0.000) ***	(0.000) ***	(0.000) ***
Cash(t-1)		1.120	1.294	1.199	0.802
		(0.571)	(0.303)	(0.453)	(0.531)
Inv(t-1)		0.982	1.052	1.008	1.173
		(0.888)	(0.718)	(0.953)	(0.382)
TangFA(t-1)		1.253	1.224	1.119	1.188
		(0.038) **	(0.101)	(0.351)	(0.330)
ROA(t-1)		0.730	0.574	0.582	1.217
		(0.332)	(0.150)	(0.161)	(0.703)
Industry dummies	Yes	Yes	Yes	Yes	Yes
Country dummies	Yes	Yes	Yes	No	Yes
Pseudo R^2	0.026	0.028	0.034	0.031	0.044
Number of obs.	9,215	8,825	7,040	7,040	3,651

Table III. The determinants of the probability of substitution by country

This table reports results from the logit regression analyses where SI_{it} is regressed on the Z-score, the different stages of the crisis, cash holdings, tangible fixed assets, inventories, firm size, ROA, and a set of industry dummies for each country separately. This regression analysis informs how the explanatory variables increase or decrease the probability of substitution between short term bank credit and accounts payable. We report the odds ratios with the *p*-values in parentheses for each explanatory variable. ***, **, * indicate coefficients that are statistically significant at the 1%, 5%, and 10% levels, respectively, using robust standard errors clustered at the firm level. All regressions in this table are conducted for external finance-dependent firms (EFD firms), as proposed by Rajan and Zingales (1998).

	(1)		(2)		(3)		(4)		(5)	
	France		Germany		Italy		Spain		U.K.	
Z(t-1)	1.184		1.082		1.156		1.087		1.099	
	(0.000)	***	(0.059)	*	(0.000)	***	(0.000)	***	(0.000)	***
D_Crisis1	0.762		0.746		0.749		0.309		0.774	
	(0.000)	***	(0.082)	*	(0.000)	***	(0.000)	***	(0.000)	***
D_Crisis2	0.576		0.449		0.635		0.480		0.274	
	(0.000)	***	(0.000)	***	(0.000)	***	(0.000)	***	(0.000)	***
D_Aftermath	0.956		1.132		1.121		0.994			
	(0.157)		(0.374)		(0.000)	***	(0.906)			
Size(t-1)	0.981		0.990		0.964		0.919		0.991	
	(0.099)	*	(0.794)		(0.000)	***	(0.005)	***	(0.405)	
Cash(t-1)	0.708		1.922		0.639		0.628		1.060	
	(0.000)	***	(0.249)		(0.000)	***	(0.109)		(0.548)	
Inv(t-1)	0.859		0.855		0.949		0.622		1.127	
	(0.023)	**	(0.644)		(0.161)		(0.000)	***	(0.187)	
TangFA(t-1)	1.268		1.617		1.219		1.142		1.245	
	(0.002)	***	(0.111)		(0.000)	***	(0.191)		(0.001)	***
ROA(t-1)	0.685		1.074		0.365		0.882		0.729	
	(0.009)	***	(0.931)		(0.000)	***	(0.704)		(0.027)	**
Industry dummies	Yes		Yes		Yes		Yes		Yes	
Pseudo R^2	0.014		0.028		0.012		0.043		0.047	
Number of obs.	31,741		1,603		81,367		12,184		20,243	

Table IV. The determinants of the probability of partial or perfect substitution

This table reports results from the multinomial regression analyses where the three-outcome variable $SI3_{it}$ is regressed on the Z-score, the different stages of the crisis, cash holdings, tangible fixed assets, inventories, firm size, ROA and a set of country and industry dummies. This regression analysis informs how the explanatory variables increase or decrease the probability of having partial or perfect substitution between short term bank credit and accounts payable relative to the base category. We report the odds ratios with the *p*-values in parentheses for each explanatory variable. ***, **, * indicate coefficients that are statistically significant at the 1%, 5%, and 10% levels, respectively, using robust standard errors clustered within firms.

	(1)		(2)		(3)		(4)	
	Full Sample		Full Sample		EFD Firms		EFD Ind.	
Dep. Var:	partial	perfect	partial	perfect	partial	perfect	partial	perfect
Z(t-1)	1.001	1.161 ***	1.009	1.166 ***	1.027	1.174 ***	0.991	1.151 ***
	(0.961)	(0.000)	(0.739)	(0.000)	(0.372)	(0.000)	(0.851)	(0.002)
D_Crisis1	0.741 ***	0.710 ***	0.750 ***	0.709 ***	0.726 ***	0.640 ***	0.616 ***	0.560 ***
	(0.002)	(0.000)	(0.004)	(0.000)	(0.005)	(0.000)	(0.003)	(0.000)
D_Crisis2	0.548 ***	0.427 ***	0.558 ***	0.429 ***	0.541 ***	0.371 ***	0.596 ***	0.418 ***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.001)	(0.000)
D_Aftermath	1.231 **	0.972	1.248 ***	0.960	1.228 **	0.871	1.399 **	0.805 *
	(0.013)	(0.719)	(0.009)	(0.610)	(0.034)	(0.118)	(0.016)	(0.087)
Size(t-1)			0.937 *	0.872 ***	0.921 *	0.841 ***	0.713 ***	0.572 ***
			(0.095)	(0.000)	(0.065)	(0.000)	(0.006)	(0.000)
Cash(t-1)			0.621 *	1.739 **	0.709	2.039 **	0.546	1.219
			(0.099)	(0.019)	(0.314)	(0.016)	(0.194)	(0.671)
Inv(t-1)			1.308 *	0.735 *	1.373 *	0.801	1.546 *	0.861
			(0.099)	(0.066)	(0.079)	(0.225)	(0.055)	(0.540)
TangFA(t-1)			1.236	1.241	1.185	1.236	1.332	1.054
			(0.129)	(0.128)	(0.291)	(0.185)	(0.205)	(0.821)
ROA(t-1)			1.102	0.507 *	0.721	0.493	2.492	0.536
			(0.817)	(0.084)	(0.501)	(0.137)	(0.142)	(0.345)
Industry dummies	Yes		Yes		Yes		Yes	
Country dummies	Yes		Yes		Yes		Yes	
Pseudo R^2	0.039		0.041		0.049		0.057	
Number of obs.	9,216		8,827		7,045		3,654	

Table V. The impact of credit quality during the financial crisis

This table reports results from the logit regression analyses where SI_{it} is regressed on the Z-score, a dummy indicator for the first and the second stage of the crisis, the interaction terms between the previous two, a vector of control variables (cash holdings, tangible fixed assets, inventories, firm size, and ROA), and a set of country and industry dummies. This regression analysis informs how the explanatory variables increase or decrease the probability of substitution between short-term bank credit and accounts payable.We report the Odds-ratios with the *p*-values in parentheses for each explanatory variable. ***, **, * indicate coefficients that are statistically significant at the 1%, 5%, and 10% levels, respectively, using robust standard errors clustered at the firm level.

	(1)	(2)	(3)
	Full Sample	EFD Firms	EFD Ind.
Z(t-1)	1.125	1.139	1.119
	(0.000) ***	(0.000) ***	(0.007) ***
D_Crisis1	0.900	0.810	0.837
	(0.424)	(0.149)	(0.393)
D_Crisis2	0.565	0.587	0.597
	(0.000) ***	(0.000) ***	(0.009) ***
*D_Crisis1*Z(t-1)*	0.906	0.930	0.847
	(0.018) **	(0.126)	(0.028) **
*D_Crisis2*Z(t-1)*	0.928	0.899	0.909
	(0.070) *	(0.022) **	(0.186)
Control Variables	Yes	Yes	Yes
Industry dummies	Yes	Yes	Yes
Country dummies	Yes	Yes	Yes
Pseudo R^2	0.028	0.034	0.045
Number of obs.	8,825	7,040	3,651

Table VI. The interaction of credit quality and financial constraints

This table reports results from the logit regression analyses where SI_{it} is regressed on the interaction terms between the Z-score and the quintile dummies for the KZ index, the quintile dummies for the KZ index separately, a vector of control variables (the different stages of the crisis, cash holdings, tangible fixed assets, inventories, firm size and ROA) and a set of country and industry dummies. This regression analysis indicates how the explanatory variables increase or decrease the probability of substitution between short term bank credit and trade credit. We report the odds ratios with the *p*-values in parentheses for each explanatory variable. ***, **, * indicate coefficients that are statistically significant at the 1%, 5%, and 10% levels respectively, using robust standard errors clustered within firms.

	(1)	(2)
	Full Sample	EFD Firms
Z(t-1)	1.155 (0.000) ***	1.166 (0.000) ***
*KZ_Q2*Z(t-1)*	1.036 (0.533)	1.080 (0.249)
*KZ_Q3*Z(t-1)*	1.096 (0.193)	1.151 (0.072) *
*KZ_Q4*Z(t-1)*	1.163 (0.040) **	1.238 (0.009) ***
*KZ_Q5*Z(t-1)*	0.953 (0.430)	0.975 (0.707)
Control Variables	Yes	Yes
KZ_Q Dummies	Yes	Yes
Industry dummies	Yes	Yes
Country dummies	Yes	Yes
Pseudo R^2	0.031	0.040
Number of obs.	8,561	6,822

Table VII. The determinants of the probability of substitution for total debt

This table reports results from the logit regression analyses where SI_{it}^{total} is regressed on the Z-score, the different stages of the crisis, cash holdings, tangible fixed assets, inventories, firm size, ROA and a set of country and industry dummies. This regression analysis informs how the explanatory variables increase or decrease the probability of substitution between total bank credit and accounts payable. We report the odds ratios with the *p*-values in parentheses for each explanatory variable. ***, **, * indicate coefficients that are statistically significant at the 1%, 5%, and 10% levels, respectively, using robust standard errors clustered within firms.

	(1)	(2)	(3)	(4)
	Full Sample	Full Sample	EFD Firms	EFD Firms demeaned
Z(t-1)	1.060 (0.000) ***	1.064 (0.001) ***	1.077 (0.001) ***	1.073 (0.001) ***
D_Crisis1	0.689 (0.000) ***	0.693 (0.000) ***	0.660 (0.000) ***	0.625 (0.000) ***
D_Crisis2	0.485 (0.000) ***	0.490 (0.000) ***	0.448 (0.000) ***	0.419 (0.000) ***
D_Aftermath	1.014 (0.825)	1.007 (0.914)	0.968 (0.652)	0.896 (0.117)
Size(t-1)		0.914 (0.001) ***	0.877 (0.000) ***	0.878 (0.000) ***
Cash(t-1)		1.062 (0.751)	1.156 (0.537)	0.129 (0.596)
Inv(t-1)		0.835 (0.154)	0.899 (0.447)	0.868 (0.302)
TangFA(t-1)		1.177 (0.106)	1.162 (0.196)	1.069 (0.558)
ROA(t-1)		0.911 (0.770)	0.773 (0.504)	0.798 (0.557)
Industry dummies	Yes	Yes	Yes	Yes
Country dummies	Yes	Yes	Yes	No
Pseudo R^2	0.025	0.025	0.032	0.029
Number of obs.	10,060	9,668	7,637	7,637

Table VIII. Analysis on the probability of substitution based on SI4

This table reports results from the multinomial regression analyses where the four-outcome variable $SI4_{it}$ is regressed on the Z-score, the different stages of the crisis, cash holdings, tangible fixed assets, inventories, firm size, ROA and a set of country and industry dummies. This regression analysis informs how the explanatory variables increase or decrease the probability of having a substitution relation (SI=2 or SI=3) or a positive complementary relation (SI=4) relative to having a negative complementary relation (SI=1). We report the odds ratios with the *p*-values in parentheses for each explanatory variable. ***, **, * indicate coefficients that are statistically significant at the 1%, 5%, and 10% levels, respectively, using robust standard errors clustered within firms.

	(1)						(2)					
	Full Sample						EFD Firms					
Dep. Var:	SI=2		SI=3		SI=4		SI=2		SI=3		SI=4	
Z(t-1)	1.084		0.897		0.965		1.097		0.889		0.966	
	(0.000)	***	(0.000)	***	(0.057)	*	(0.000)	***	(0.000)	***	(0.119)	
D_Crisis1	0.734		1.011		0.698		0.685		0.992		0.648	
	(0.000)	***	(0.883)		(0.000)	***	(0.000)	***	(0.926)		(0.000)	***
D_Crisis2	0.488		0.768		0.374		0.443		0.770		0.321	
	(0.000)	***	(0.000)	***	(0.000)	***	(0.000)	***	(0.001)	***	(0.000)	***
Control Variables	Yes						Yes					
Industry dummies	Yes						Yes					
Country Dummies	Yes						Yes					
Pseudo R^2	0.022						0.026					
Number of obs.	18,802						15,070					

Table IX. Substitution controlling for contemporaneous shocks to bank credit

This table reports results from the logit regression analyses where SI_{it} (with an aggregate negative shock to bank credit in year t-1 and year t) is regressed on the Z-score, the different stages of the crisis, cash holdings, tangible fixed assets, inventories, firm size, ROA and a set of country and industry dummies. This regression analysis informs how the explanatory variables increase or decrease the probability of substitution of short term bank credit for accounts payable. We report the odds ratios with the *p*-values in parentheses for each explanatory variable. ***, **, * indicate coefficients that are statistically significant at the 1%, 5%, and 10% levels, respectively, using robust standard errors clustered at the firm level.

	(1)		(2)	
	Full Sample		EFD Firms	
Z(t-1)	1.094		1.101	
	(0.000)	***	(0.000)	***
D_Crisis1	0.730		0.679	
	(0.001)	***	(0.001)	***
D_Crisis2	0.481		0.447	
	(0.000)	***	(0.000)	***
Control Variables	Yes		Yes	
Industry dummies	Yes		Yes	
Country dummies	Yes		Yes	
Pseudo R^2	0.029		0.034	
Number of obs.	6,346		5,041	

Table X. Stratified Random Sampling

This table reports the average estimation results from the logit regressions in which we regress SI_{it} on the Z-score, a dummy variable for the first and the second stage of the crisis, the interaction terms between the previous two, a vector of control variables (cash holdings, tangible fixed assets, inventories, firm size, and ROA), and a set of country and industry dummies for 100 randomly drawn stratified samples. For each country, the firms are divided in size quintiles and within each quintile we have drawn a fixed number of firms. The sample includes firms that are external finance-dependent (EFD firms), as proposed by Rajan and Zingales (1998) or only external finance-dependent firms conditional on non-negative change in total assets. The table reports the mean odds ratios, the mean *p*-values, and the mean pseudo R-squares for the 100 regression analysis. ***, **, * indicate coefficients that are statistically significant at the 1%, 5%, and 10% levels, respectively, using robust standard errors clustered at the firm level.

	Replication Table 2				Replication Table 5			
	(1)		(2)		(3)		(4)	
	EFD Firms		EFD Firms, dTA≥0		EFD Firms		EFD Firms, dTA≥0	
Z(t-1)	1.110		1.092		1.119		1.058	
	(0.000)	***	(0.013)	**	(0.000)	***	(0.107)	
D_Crisis1	0.681		0.676		0.682		0.417	
	(0.000)	***	(0.001)	***	(0.023)	**	(0.001)	***
D_Crisis2	0.477		0.740		0.549		0.725	
	(0.000)	***	(0.004)	***	(0.000)	***	(0.106)	
*D_Crisis1*Z(t-1)*					1.000		1.177	
					(0.986)		(0.022)	**
*D_Crisis2*Z(t-1)*					0.961		1.018	
					(0.378)		(0.421)	
Control Variables	Yes		Yes		Yes		Yes	
Industry Dummies	Yes		Yes		Yes		Yes	
Country Dummies	Yes		Yes		Yes		Yes	
Pseudo R2	0.029		0.040		0.029		0.040	

Table XI. Analysis with matched bank-firm data

This table reports the logit regression with matched bank-firm data from Spain. We regress SI_{it} on a dummy variable for the financial crisis (*Crisis*), a dummy for banks that were bailed out (*Bailout*), the interaction variable *Bailout*Crisis*, time-varying firm controls, time-varying bank controls, and industry fixed effects. The variable *NumberRel* indicates the number of bank relationships per firm and year. In column 4 we replace the *Bailout* dummy with the *SavingsBank* dummy. We report odds ratios and *p*-values in parentheses. All firms are dependent on external finance following Rajan and Zingales (1998). ***, **, * indicate coefficients that are statistically significant at the 1%, 5%, and 10% levels, respectively, using robust standard errors clustered at the bank-firm level.

	(1)	(2)	(3)	(4)
Crisis	0.717	0.857	0.860	0.879
	(0.000) ***	(0.000) ***	(0.000) ***	(0.000) ***
Bailout	1.080 ***	1.117	1.096 **	
	(0.018)	(0.001) ***	(0.010)	
*Bailout*Crisis*	0.841	0.842	0.837	
	(0.000) ***	(0.000) ***	(0.000) ***	
SavingsBank				1.122 ***
				(0.001)
*SavingsBank*Crisis*				0.828
				(0.000) ***
Z(t-1)		1.329	1.335	1.336
		(0.000) ***	(0.000) ***	(0.000) ***
Size(t-1)		1.459	1.452	1.452
		(0.000) ***	(0.000) ***	(0.000) ***
Cash(t-1)		0.356	0.366	0.366
		(0.000) ***	(0.000) ***	(0.000) ***
TangFA(t-1)		1.087	1.097	1.099
		(0.088) *	(0.078) *	(0.075) *
Inv(t-1)		0.550	0.544	0.545
		(0.000) ***	(0.000) ***	(0.000) ***
ROA(t-1)		0.292	0.292	0.294
		(0.000) ***	(0.000) ***	(0.000) ***
NumberRel			1.052	1.052
			(0.000) ***	(0.000) ***
Bank_ROA			0.061	0.095
			(0.399)	(0.476)
Bank_eqta			1.402	1.388
			(0.534)	(0.557)
Bank_depta			1.180	1.065
			(0.175)	(0.527)
Industry Dummies	Yes	Yes	Yes	Yes
Pseudo R^2	0.011	0.028	0.029	0.029
Number of obs.	59,534	51,304	51,304	51,304

Figure 1. The substitution indicator over time

This figure shows the substitution indicator SI_{it} by country and over time. The years are shown on the x-axis, while the fractions of substitution relationships (i.e., $SI_{it} = 1$) are shown on the y-axis.

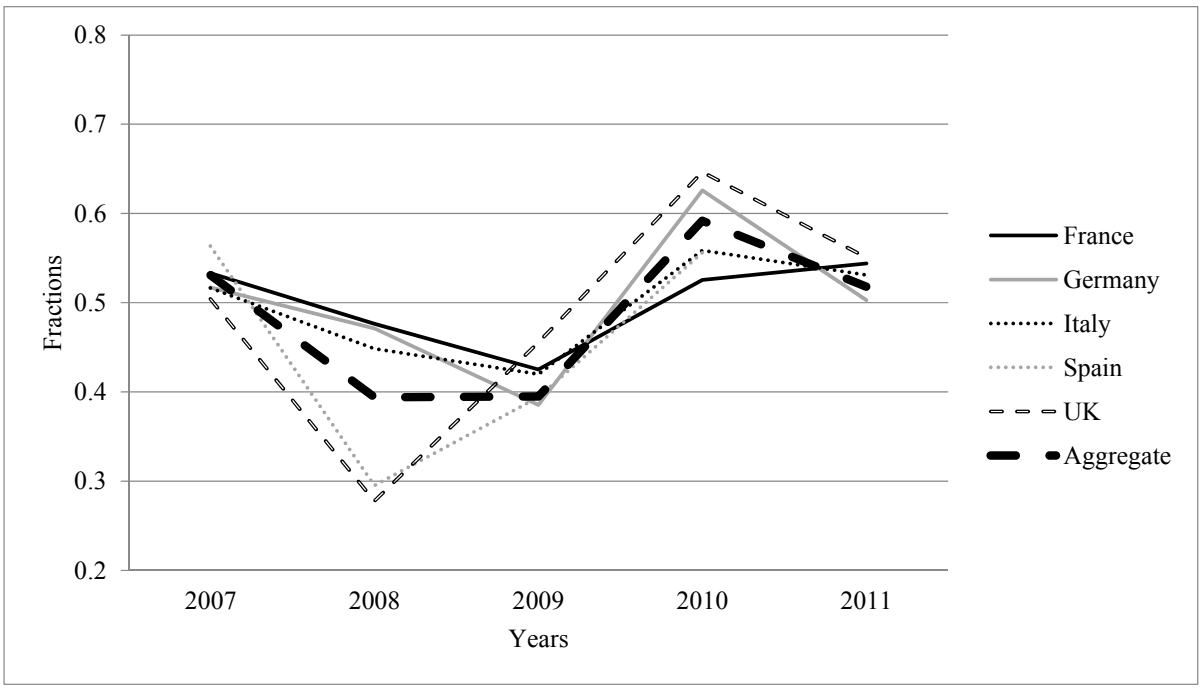

Figure 2. Odds ratios of the interaction terms between Z-score and KZ index quintiles

This figure presents the odds ratios for the interacted *Z-Score* (*t*-1) and KZ quintile dummies by country on the y-axis and the KZ index quintile dummies on the x-axis. The KZ quintiles are computed separately for each country.