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Trade Credit and Pricing: An Empirical Evaluation*

Niklas Amberg [†] Tor Jacobson [‡] Erik von Schedvin [§]

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Abstract

We empirically investigate the proposition that firms charge premia on cash prices in transactions involving trade credit. Using a comprehensive Swedish panel dataset on product-level transaction prices and firm-characteristics, we relate trade credit issuance to price setting. In a recession characterized by tightened credit conditions, we find that prices increase significantly more on products sold by firms issuing more trade credit, in response to higher opportunity costs of liquidity and counterparty risks. Our results thus demonstrate the importance of trade credit for price setting and show that trade credit issuance induces a channel through which financial frictions affect prices.

Keywords: Trade credit; prices; inflation; liquidity; counterparty risk.

JEL: E31; E32; D22; G30; L11.

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1 Introduction

The notion that sellers may charge substantial interest rates when extending trade credit is based on the discount structure in two-part terms contracts (see, e.g., Petersen and Rajan, 1994, 1997) and has reputed this a costly source of short-term funding. Yet, trade credit is commonly issued under net terms contracting, for which no price discount is offered, nor is an explicit interest rate specified.[1] Thus, the nature of trade credit pricing is ambiguous and it remains an open question whether trade credit is indeed expensive, or not.

We conjecture that sellers, under net terms contracting, price trade credit by adding an interest rate premium to the price of the good. That is, sellers incorporate the transaction time value of money into their pricing decisions. Accordingly, the premium should increase in the maturity of the trade credit contract, the seller's cost of liquidity, and the buyer's default risk. Since trade credit is an abundant feature of inter-firm trade, shifts in the trade credit price premium—caused, for instance, by the credit supply contractions and rises in counterparty risks typically observed in periods of financial distress—could have large impacts on prices and make for an important channel through which financial frictions interact with prices.[2] However, there is to this date no empirical work that has examined the role of firms' trade credit issuance for their pricing decisions, possibly reflecting a scarcity of firm-specific product-level price data.

We aim to contribute towards a better understanding of firms' trade credit pric-

[1] For US firms, Ng, Smith and Smith (1999) find that 25.5 percent of the firms in their sample mainly offer two-part contracts, while Giannetti, Burkart and Ellingsen (2011) show that only 31 percent of the firms in the 1998 National Survey of Small Business Finances had been offered early payment discounts from their main seller. Using more recent data from 2005, Klapper, Laeven and Rajan (2012) find that only 13 percent of the contracts extended to a sample of large US and European buyers included early payment discounts. In Sweden, which is our empirical setting, net terms contracts are predominantly used.

[2] In Sweden, 97 percent of the business-to-business transactions involve trade credit (Pärlhem, 2016). Moreover, Jacobson and von Schedvin (2015) show that the average amount of accounts receivable and payable, scaled by assets, are 16 and 11 percent for Swedish firms. Similar reliance on trade credit financing prevails across countries. For instance, Rajan and Zingales (1995) show that the corresponding numbers for a sample of US firms are 18 and 15 percent, and Berger and Udell (1998) show that trade credit provides 31 percent of debt financing to US SMEs, nearly as much as commercial banks (37 percent).

ing, and thereby of product pricing in general. To this end, we make use of a dataset comprising product-level data on prices and quantities for all Swedish manufacturing firms; firm-level accounting data; and loan-level data covering all loans extended by the four major Swedish banks to Swedish corporations. These data allow us to relate firm-specific product-level inflation rates to trade credit issuance at the firm-level—which is unique in this empirical context—while carefully assessing the robustness of our results and validate the plausibility of our identifying assumptions. More specifically, our empirical design is geared to assess the importance of trade credit issuance for price-setting in the 2008–09 recession in Sweden. The recession—characterized by a severe credit crunch, as well as a sharp downturn in the real economy—led to increases in the costs of liquidity and counterparty risks facing firms and is of key importance for our identification that exploits differences in trade credit issuance across firms.

Our main finding can be summarized as follows. Firms that issued trade credit with longer average maturities, relative to firms that issued shorter average maturities, increased their prices significantly more in the 2008–09 recession. The estimated increase in the annualized trade credit interest rate over the crisis episode amounts to 41.2 percentage points. In terms of price adjustments, this trade credit interest rate implies that a maturity difference of 20 days (approximately equivalent to a one-standard deviation in the maturity distribution) lead to a relative price adjustment of 2.3 percentage points. This effect magnitude can be put into perspective by the following considerations. Firstly, our hypothesis is that the increase in trade credit interest rates in part reflects a tightening in external financing and the associated rise in the opportunity costs of liquidity facing sellers, and in part a rise in credit risk premia, both of which may have been substantial in the crisis, see Whited (1992) for the former aspect and Berndt, Douglas, Duffie, and Ferguson (2018) for the latter. Secondly, the documented effect is close in magnitude to the annualized implicit interest rate for the well-known "2/10 net 30" two-part terms contract, amounting to 44.6 percent (see, e.g., Giannetti, Burkart and Ellingsen, 2011). Thirdly, another useful benchmark is the usage of factoring services, i.e., sales of sellers' invoices to a third party priced at a discounted nominal invoice value.[3] Factoring discounts in Sweden are currently in the range of 2 to 5 percent, which in the case of a net 30 contract corresponds to implicit annualized interest rates of 24.6 to 62.4 percent. Hence, our estimated increase in trade credit interest rates is substantial but quite reasonable when considering shifts in the underlying determinants of trade credit interest rates, the implicit interest rates on two-part terms contracts, and the costs associated with factoring.

The assumptions underlying identification are validated in several ways. Firstly, there is no significant divergence in pre-treatment pricing trends between low and high trade credit firms, which speaks against the documented effects being present in the absence of the 2008–09 recession. Secondly, we show that the impact of trade credit issuance on prices remains when we control for a broad set of factors that previous research has shown to be important determinants of firms' price setting, which relaxes the concern that our results are outcomes of spurious correlations. For example, in line with the empirical results presented by Chevalier and Scharfstein (1996), Campello (2003), and Gilchrist et al. (2017), we find a direct impact of firms' pre-crisis leverage and liquidity positions on their pricing during the crisis; which is consistent with the view that liquidity constraints may give rise to countercyclical price mark-ups. Nevertheless, the effect of trade credit persists once these alternative factors are accounted for, indicating that we document a complementary channel through which financial frictions operate. Thirdly, by exploring cross-sectional heterogeneity we show that the impact of trade credit issuance on prices is primarily observed for liquidity-constrained firms, as measured by cash holdings and access to bank lines of credit. Moreover, effects are more pronounced for firms that experienced a larger deterioration in the credit quality of their customer bases. These cross-sectional results support our main conjecture that the opportunity cost of liquidity and counterparty risk are key drivers of the documented effects. Finally, in our empirical analysis we also evaluate an alternative exposition for a positive relationship between trade credit maturities and prices arising only due to shifts in buyers' demand for trade credit; with no role for sellers' supply of credit. Un-

[3] Only a small fraction, around 3 percent, of Swedish firms use factoring services (Pärlhem, 2016).

der this scenario, quantities should increase in the crisis period for goods sold by sellers issuing longer trade credit maturities. However, we find a no relationship between trade credit maturities and changes in quantities; pointing to the relevance of a supply-side role of trade credit according to our main conjecture of a trade credit pricing premium.

The existing literature on the pricing of trade credit is fairly limited and has primarily concerned implications of two-part terms contracts, providing ambiguous evidence. On the one hand, Ng, Smith and Smith (1999) provide survey evidence suggesting that credit terms under two-part contracts are persistent over time and unrelated to shifts in the sellers' cost of funds, and Petersen and Rajan (1994) document that interest rates on trade credit typically are unrelated to the quality of the buyer. On the other hand, in a more recent paper, Klapper, Laeven and Rajan (2012) show that early payment discounts primarily appear to be used as a risk management tool by small exposed sellers attempting to incentivise early payments from low-quality buyers, in order to avoid defaults and subsequent credit losses. The findings in our paper thus contribute to the trade credit literature by documenting that sellers' valuation of liquidity and counterparty risks—under net terms contracting—is implicitly priced by adding a premium to the product price. In our view, due to the widespread use of net terms contracts, these findings are of key importance towards a better understanding of trade credit pricing.[4]

Furthermore, by issuing trade credit, firms function as financial intermediaries in terms of screening, monitoring, and lending to counterparties. Therefore our results also relate to the extensive literature on the bank lending channel. For example, during the recent financial crisis, US banks faced sharp increases in funding costs, caused primarily by exposures to subprime assets and the breakdown of the interbank market (see, e.g., Cornett et al., 2011; Chodorow-Reich, 2014). In re-

[4]Even in settings where sellers make use of two-part terms contracts, it could well be that pricing is partly done through a premium on the cash price, and not entirely through the early payment discount. This may explain previous findings, showing mixed evidence for counterparty risks, and no role for sellers' liquidity costs. Also, and intuitively, two-part contract sellers' inclination towards pricing by premium should increase with the length of the discount period; Cuñat and Garcia-Appendini (2012) point to the existence of "8/30 net 50"-contracts, for which the discount period is 30 days, making the first part of this version of two-part contracts equivalent to a net 30 contract.

sponse, banks tended to increase loan interest rates (Santos, 2010) and in this dimension our findings align with the actions reported for US banks. But, contrary to our results—where trade credit issuance is somewhat reduced over the crisis—evidence for the bank lending channel show that adjustments of loan quantities are of first order importance (see, e.g., Ivashina and Scharfstein, 2010). An intuitive explanation for this difference in crisis response between trade credit and bank credit is that higher bank lending interest rates may exacerbate adverse selection and moral hazard facing banks, making quantities more tractable to adjust than prices (Stiglitz and Weiss, 1981). Since trade credit is goods-specific, with sellers servicing a sticky customer base and goods are harder than cash to divert, adverse selection and moral hazard may be less of a concern in trade credit issuance, thus making adjustments in prices more feasible.

Finally, as noted above, recent research on the links between firm-characteristics and price setting has demonstrated that due to capital market imperfections, firms' liquidity and leverage positions make for important determinants of movements in price mark-ups over the business cycle (see Chevalier and Scharfstein, 1996; Campello, 2003; and Gilchrist, Schoenle, Sim, and Zakrajšek, 2017).[5] Our findings contribute to this literature by unravelling an additional channel at the firm-level through which seller and buyer characteristics affect price-setting.

The rest of the paper is organized as follows. The next section presents a conceptual framework that outlines the link between trade credit and pricing. Section 3 describes the 2008–09 recession in Sweden, details our data resources, and provides some descriptive statistics. The empirical framework is presented in Section 4 and the results are presented in Section 5. Section 6 concludes.

[5]These findings are consistent with theoretical contributions by Gottfries (1991) and Chevalier and Scharfstein (1996), proposing that liquidity constrained firms operating in customer markets—markets in which the customer base is sticky—increase prices to enhance profits during economic downturns, at the expense of future market shares.

2 Conceptual Framework

Trade credit arises in inter-firm trade, when sellers allow buyers to fulfil payment at a date later than that of delivery. Trade credit contracts are formulated in either net terms, or in two-part terms, where the former simply specifies a due date, and the latter also includes a price discount should the buyer pay before the due date. As argued in the previous section, pricing of trade credit applies to both types of contract, which is at odds with the notion that trade credit pricing only occurs in two-part contracts in the case when the buyers choose to disregard the price discount and make use of the full payment period. In our Swedish setting, the net days-type of contract is the currently prevailing—not to say the only one. We will therefore let the conceptual relationship between trade credit and firms' pricing behaviour outlined below assume the net days-type of trade credit arrangement, but with no loss of generality in the sense that pricing of trade credit for two-part contracts could well be similar, except for the explicit credit cost arising in discounts foregone.

Thus, towards a conceptual framework to guide the subsequent empirical analysis, our point of departure is the standard formulation of the firm's price-setting problem. The optimal price for product p sold by firm i, $P_{i,p}$, is equal to the product of the firm's marginal cost for producing p, $MC_{i,p}$, and a mark-up, $\mu_{i,p}$, that depends on the firm's price-setting power in the product market:

$$P_{i,p} = \mu_{i,p} \cdot MC_{i,p}. \tag{1}$$

This characterization of the price-setting problem neglects one salient aspect, however, namely that inter-firm transactions ever so often involve trade credit.

Since lending is associated with costs—most importantly due to funding and to credit risk exposure—prices charged in trade credit transactions likely surpass prices charged in cash transactions. Schwartz (1974) highlights this trade credit feature of price-setting and posits that firms add a trade credit premium to the cash price, determined by contracted loan maturity and an implicit interest rate.[6] Our

[6]More recently, Udell (2015) discusses what he labels the trade credit pricing puzzle, and notes that: "the 'all-in' price – that which matters – must incorporate both the price of the product as well

conceptual framework—intended to support the subsequent empirical analyses—rests on the relationship proposed by Schwartz and we focus on its implications for the link between firms' trade credit issuance and pricing decisions.

To formalize, let $P_{i,p}^{C}$ denote the cash price, corresponding to the price in Eq. (1), and let $P_{i,p}^{T}$ denote the trade credit price. $P_{i,p}^{T}$ can then be expressed as a function of said cash price, the maturity, and an interest rate:

$$P_{i,p}^{T} = P_{i,p}^{C} \cdot e^{r_{i,p} \cdot \tau_{i,p}}, \tag{2}$$

where $r_{i,p}$ is the implicit annual interest rate charged by the seller and $\tau_{i,p}$ is the maturity of the trade credit contract, in number of net days divided by 365. The interest rate and maturity may well vary across transactions; the parameters $r_{i,p}$ and $\tau_{i,p}$ should therefore be interpreted as averages across all sales of product p by firm i.

From Eqs. (1) and (2), we can derive the firm-product inflation rate:

$$\begin{aligned}
\pi_{i,p,t}^{T} &= \ln P_{i,p,t}^{T} - \ln P_{i,p,t-1}^{T} \\
&= \Delta \ln \mu_{i,p,t} + \Delta \ln MC_{i,p,t} + \Delta \left(r_{i,p,t} \cdot \tau_{i,p,t} \right) \\
&= \Delta \ln \mu_{i,p,t} + \Delta \ln MC_{i,p,t} + \left(\tau_{i,p,t-1} \cdot \Delta r_{i,p,t} + r_{i,p,t-1} \cdot \Delta \tau_{i,p,t} \right).
\end{aligned} \tag{3}$$

If we assume that maturities are approximately constant over time, implying that $\Delta\tau \approx 0$, the firm-product inflation rate in year t is determined by the change in the mark-up, the change in marginal cost, and the product of the average trade credit maturity and the change in the implicit interest rate; all of which are allowed to vary at the firm-product level. If firms, on the contrary, can adjust maturities in trade credit contracts in response to shifts in liquidity costs and counterparty risks, they may choose to shorten maturities when such shifts occur. This would attenuate the relationship between r and π^{T}, since some part of the direct effect of an increase in r on π^{T} would be offset by the decrease in τ. We return to the issue of how changes

as the financial terms of trade credit. It is highly unlikely that there are any available data that would allow us to calculate this all-in price." Moreover, early empirical work by Hekman (1981) document the extent to which trade credit price premia matter for cross-country comparisons of product prices.

in trade credit contract maturities may affect our results in Section 4.2 below.

In line with Schwartz (1974), our hypothesis is that the trade credit interest rate r is determined primarily by two factors: (i) the seller's cost of liquidity and (ii) the buyer's risk of defaulting on the transaction.[7] That is, the implicit interest rate underlying trade credit is increasing in sellers' liquidity costs, as well as in their credit risk exposures, all else equal. The cost of liquidity, in turn, that applies to the issuance of trade credit is given by the opportunity cost of the marginal unit of liquidity facing the firms—commonly denoted as the shadow cost of liquidity—and is thus equal to the external funding cost for financially unconstrained firms, but higher than this for firms that experience binding liquidity constraints due to external financing rationing.[8] For counterparty risks, the Bankruptcy Code in Sweden, dictating terms for bankruptcy resolution, provides trade credit debt with junior priority status, and is thus similar to international practice, see Cuñat and Garcia-Appendini (2012) for an overview. This means that sellers have little hope of recovering claims on failed buyers' bankruptcy estates after prioritized holders of claims have been handled, see Thorburn (2000).

It is unlikely that firms can set higher prices than their competitors for prolonged periods, since at some point buyers will overcome switching frictions and turn to sellers offering lower prices. Permanent differences in liquidity costs across firms should therefore not be reflected in corresponding price differences across firms. However, if a firm operates in a customer market, i.e., a market in which the customer base is sticky—for instance because of costly switching (Klemperer, 1987), costly search (Hall, 2008), or idiosyncratic preferences (Bronnenberg, Dubé

[7]Antitrust legislation may to some extent limit firms from engaging in price discrimination. Similarly to the US setting, EU Competition Law (Article 82(c) of the EC Treaty) dictates what Swedish firms can and cannot do in terms of setting differential prices across buyers, and under what circumstances. Whereas it is true that firms cannot fully customize prices—although Article 82(c) has quite strict and precise requirements for making a case of price discrimination—free pricing is not a prerequisite for our analysis, neither conceptually, nor empirically. Indeed, the average interest rate and maturity parameters in Eq. (2) are consistent with firms setting standardized prices equal for all their buyers on the basis of the opportunity costs of liquidity they face and the average counterparty risks in the pool of buyers they service.

[8]See Whited (1992) for a detailed outline of the role of financial constraints and the shadow cost of external financing. She documents that the marginal value of one extra unit of liquidity is substantially higher for financially constrained firms, as compared with unconstrained firms that have bond ratings and access to market funding.

and Gentzkow, 2012)—then its prices may differ from competitors' prices in the short-run. Phelps and Winter (1970) and Bils (1989) show that an important feature of price setting in customer markets is the trade-off between maximizing short-term revenue by increasing prices and building a future customer base by lowering prices. This suggests that it may be optimal for firms to pass on temporary variation in liquidity costs to buyers in trade credit transactions, in particular the sharp but temporary increases in cost of liquidity that typically occur during financial crises.

The same reasoning is not necessarily true for the credit risk component, since any potential seller faces the same credit risk when extending trade credit to a given buyer and therefore will require the same actuarially fair compensation for bearing this risk—abstracting from differences in credit risk attributable to the terms in the trade credit contract. Hence, the credit risk component of the implicit interest rate may well reflect both temporary and permanent differences in counterparty risk faced by different sellers.[9]

We close this section in consideration of one alternative explanation for a positive relationship between trade credit issuance and prices during times of credit tightening reflecting shifts in buyers' demand for trade credit. In this narrative, firms that issue longer maturities would face a higher demand for their goods on account of buyers' increased demand of credit when external credit becomes scarce, which in turn would push up these sellers' prices relative to firms that provide short maturities. However, an outward shift in demand can be expected to have a positive impact on quantities sold for firms that provide long maturities relative to short maturity firms. Conditional on an outward shift, the only possibility for unchanged, or a relative decline, in long maturity firms' quantities is if the demand effect is offset by a price increase stemming from the supply-side; caused by, for instance, the factors proposed above. By way of including quantities sold in the empirical analysis,

[9]The aggregate corporate bankruptcy rate in Sweden almost doubled, as the crisis episode unfolded. However, it is conceivable that shifts in corporate credit risk premia were even more dramatic. Recently Berndt et al. (2018) found that risk premia, measured as differences between credit default swap rates and expected default loss rates, for US public firms peaked in the global financial crisis of 2008–09. Costs for credit insurance per unit expected loss increased by a factor 10. In addition, Berndt et al. find substantial cross-sectional differences in credit risk premia and also note that premia are increased by market illiquidity.

this alternative interpretation can be evaluated and we outline the test below.

3 Setting, Data, and Descriptive Statistics

3.1 The 2008–09 recession in Sweden

We exploit the 2008–09 recession in Sweden to test the hypothesis that trade credit prices include a price premium determined by an implicit interest rate reflecting sellers' liquidity costs and credit risk exposures. The 2008–09 recession is well suited for this purpose, in featuring a sharp downturn in the real economy as well as severe distress in the banking sector—and since the origin of both lay in external shocks hitting the Swedish economy in the wake of the global financial crisis.[10]

The banking sector distress was largely due to two external shocks. Firstly, the collapse of international financial markets following the outbreak of the subprime crisis in the US. While Swedish banks had little direct exposure to mortgage-backed securities issued in the US, the Swedish banking sector is highly dependent on external wholesale funding and therefore sensitive to conditions on international financial markets. Secondly, the severe economic crisis in the Baltic countries caused large loan losses for two of Sweden's four major banks, which had expanded rapidly on the Baltic market prior to the crisis.[11] The Baltic crisis naturally affected the exposed banks more, but the unexposed banks were partly affected as well, since the problems stemming from the Baltic countries gave rise to concerns about the stability of the Swedish banking sector as a whole. These two shocks led to increased distress in the banking system, although observers' judgments differ somewhat as to the severity of the distress. According to the IMF's banking crisis database, for example, Sweden suffered a "borderline" systemic banking crisis beginning in 2008 (Laeven and Valencia, 2012), while Romer and Romer (2017), using a financial distress measure ranging from 0 to 15, classifies the level of distress in Sweden during

[10]See, e.g., Bryant, Henderson and Becker (2012) for a comprehensive discussion of the causes and consequences of the 2008–09 recession in Sweden.

[11]The Swedish bank market is dominated by four major banks, jointly accounting for around 85 percent of banking sector assets and 75 percent of corporate lending.

2008–09 as 5 on average, with a peak value of 7.

The banking sector distress quickly led to a deterioration in the credit conditions facing corporate borrowers: beginning in 2008 and continuing throughout 2009, growth in bank lending to firms fell steadily (Finansinspektionen, 2012), and many firms reported on a worsening access to external finance (Sveriges Riksbank, 2009; Konjunkturinstitutet, 2009). Meanwhile, the real economy fell into a sharp recession, with a decline in real GDP of around six percent in 2009, partly due to the domestic banking sector distress and partly due to the breakdown in international trade, which hit the export-oriented Swedish economy badly. The recession did not only hurt firms' cash flows, and thereby their liquidity positions, but also led to increased counterparty risks, not least manifested in a doubling of aggregate bankruptcy rates. Thus, the events unfolding during the 2008–09 recession increased liquidity costs as well as credit risk exposures in the corporate sector; both of which yielded a rise in r, according to the hypothesis outlined in the previous section.

3.2 Data and variable definitions

The empirical analysis in this paper is based on data from four sources, which we merge unambiguously by means of the unique identifier (*organisationsnummer*) attached to each Swedish firm. Firstly, we obtain data on prices and quantities from "Industrins varuproduktion", an annual survey conducted by Statistics Sweden comprising all manufacturing plants with at least 20 employees, as well as a sample of smaller plants. The data cover transaction prices and quantities of goods sold at the product-plant level; where products are classified using 8/9-digit CN codes.[12] Thus, for each product produced at a given plant, we observe the average transaction price (as opposed to the list price), as well as the quantity of goods sold in each year. We aggregate the price and quantity data to the firm-product level using the sales value for each product and plant as weights.

[12]These data have previously been used by Carlsson and Skans (2012). To give an idea of the granularity of the product classification, we can, for example, consider the codes 84212100 and 84212200, which refer to 'machinery and apparatus for filtering or purifying water' and 'machinery and apparatus for filtering or purifying beverages (excl. water)', respectively.

Secondly, we obtain firm-level accounting data from the database Serrano, which covers the universe of corporations in Sweden. Serrano is constructed based on data from several official sources, most importantly the Swedish Companies Registrations Office, to which all Swedish corporations are required to submit annual financial accounting statements in accordance with EU standards. Thirdly, we use a loan-level database available at Sveriges Riksbank, which covers all loans and credit lines extended by the four major Swedish banks to Swedish corporations. Finally, we obtain data on firm-level default probabilities from UC AB, the leading credit bureau in Sweden.

Our primary outcome variable is the firm-product inflation rate, defined as the log change in average transaction prices for product p, charged by firm i, between $t-1$ and t:

$$\pi_{i,p,t} = \ln P_{i,p,t} - \ln P_{i,p,t-1}.$$

We obtain 49,134 firm-product inflation rate observations, corresponding to 3,928 firms and 3,917 unique products, over the sample period 2004–2011.[13] Panel A of Figure 1 shows the distribution of annual firm-product inflation rates. Around 16 percent of the observations are located in the ± 0.5 percent interval around zero, while around half are larger than 5 percent in absolute value.

The main explanatory variable concerns firms' trade credit maturities. For want of contract-level data and the exact maturity in each trade credit contract, we use the ratio of accounts receivable to sales divided by 365:

$$\hat{\tau}_i^{07} = \frac{Rec_i^{07}}{Sales_i^{07}} \cdot \frac{1}{365}.$$

$\hat{\tau}_i^{07}$ is thus a proxy for firm i's average trade credit maturity across all its products and buyers in 2007.[14] We fix this variable to its last pre-crisis value to mitigate endogeneity concerns, but we confirm below that our results are robust to allowing τ

[13]The data contain several observations of very large price changes, which may well reflect unobserved changes in product quality. We remove such observations by truncating the inflation rate variable at the 5th and 95th percentiles.

[14]More precisely, $\hat{\tau}_i^{07}$ measures average time to payment, which may differ from contracted payment time due to either late or premature payments.

to vary over time.

We include two sets of control variables. The first set of controls, $\mathbf{X_{i,p,t}}$, consists of two variables at the firm-product level: the log change in the quantity of sales of product p by firm i between years $t-1$ and t, $\Delta Q_{i,p,t}$; and the change in unit input costs for product p produced by firm i between years $t-1$ and t, $\Delta UIC_{i,p,t}$. Unit input costs are defined as the sum of labor costs and intermediate input costs divided by physical output.[15] The second set of controls, $\mathbf{Z_{i,t-1}}$, comprises the following firm-level variables: cash and liquid assets, $Cash/Assets_{i,t-1}$; leverage, $Total\ debt/Assets_{i,t-1}$; asset tangibility, $Tangible\ assets/Assets_{i,t-1}$; cash flow, $EBITDA/Assets_{i,t-1}$; inventories, $Inventories/Sales_{i,t-1}$; and firm size, $\ln Assets_{i,t-1}$. We winsorize the explanatory variables at the 1st and 99th percentiles to reduce the influence of outliers.

3.3 Sample and descriptive statistics

Table 1 reports descriptive statistics for all variables used in the empirical analysis. The mean (median) firm-product inflation rate, reported in Panel A, is 2.8 (0.6) percent. The average value of $\hat{\tau}_i^{07}$, reported in Panel B, is 0.097, corresponding to a trade credit contract maturity of 35 days. Sixty percent of the firms have access to a non-exhausted credit line and the average size of the unused part is 4.3 percent of total assets. Panel C, finally, shows the values of the time-varying firm-characteristics. The average firm has a book value of assets of 283 million SEK and sales of 355 million SEK (roughly 44 and 55 million USD, respectively, at the exchange rate prevailing at the end of 2007). The sample thus consists primarily of medium and large firms.

[15] We do not observe labor costs and intermediate input costs at the product level, so we must resort to the following approximation when calculating $\Delta UIC_{i,p,t}$. For each plant and year, we portion out plant-level labor costs and intermediate input costs across the products produced at the plant in proportion to the total plant production share of the product. We approximate the share of each product as the physical quantity produced during the year times the firm's average unit price for the product over the entire sample period. We define the shares in this way—instead of simply using the shares of each product in total sales at the plant—in order to avoid introducing spurious positive correlation between price changes and $\Delta UIC_{i,p,t}$. We then compute the plant-product level values of $\Delta UIC_{i,p,t}$. Finally, we aggregate $\Delta UIC_{i,p,t}$ to the firm-product level using total sales for each product-plant as weights.

In Panel B of Figure 1, we show that our sample is representative of the broader economy in terms of price changes. More specifically, the figure shows that the average firm-product inflation rates in our sample tracks the changes in the aggregate producer price index for the goods-producing sector of the economy fairly closely over the entire sample period.

4 Empirical Framework

4.1 Empirical strategy

Our empirical strategy is to exploit an aggregate shock that generically increased costs of liquidity and counterparty risks, and achieve identification using the cross-sectional variation in τ, which determines the sensitivity of firms' prices to such increases, cf. Eq. (3). The identifying variation thus comes from the differences in average trade credit maturities across firms that prevailed at the time of the aggregate shock to r. Previous research has documented several factors that affect trade credit contract maturities, including financial factors (Garcia-Appendini and Montoriol-Garriga, 2013), product characteristics (Giannetti, Burkart and Ellingsen, 2011), market power (Klapper, Laeven and Rajan, 2012), and legislation (Barrot, 2016). An identifying assumption underlying our strategy—to be addressed in more detail below—is that this variation is uncorrelated with any unobserved factors that affected firms' price setting in the crisis.

The hypothesis that increases in liquidity costs and credit risk exposure cause firms to increase product prices can be tested using the following difference-in-differences specification:

$$\pi_{i,p,t} = \beta \cdot Crisis_t \cdot \hat{\tau}_i^{07} + \alpha_{i,p} + \alpha_t + \gamma \cdot \mathbf{X_{i,p,t}} + \delta \cdot \mathbf{Z_{i,t-1}} + \varepsilon_{i,p,t}. \quad (4)$$

where $\pi_{i,p,t}$ is the firm-product inflation rate; $Crisis_t$ is a dummy variable equal to one in the years 2008 and 2009 and zero otherwise; $\hat{\tau}_i^{07}$ is the average trade credit maturity for firm i in 2007; $\alpha_{i,p}$ and α_t are firm-product and year-fixed effects, respectively; and $\mathbf{X_{i,p,t}}$ and $\mathbf{Z_{i,t-1}}$ are the vectors of control variables defined in Section 3.2. The firm-product fixed effects control for potential time-invariant differences in price setting between firms with low and high trade credit issuance, respectively, while the vector $\mathbf{X_{i,p,t}}$ controls for fluctuations in demand and production costs at the firm-product level. The vector $\mathbf{Z_{i,t-1}}$, finally, controls for additional time-varying firm-characteristics that may influence price setting. Standard errors are clustered at the firm-level in all regressions.

We estimate the baseline specification for the period 2004–2011, which comprises a four-year pre-crisis period (2004–2007), the crisis period itself (2008–2009), and a two-year post-crisis period (2010–2011). The coefficient of interest is β, measuring the extent to which trade credit maturities affected firm-product inflation rates in the crisis period. More specifically, following from Eq. (3) and the assumption of persistence in τ, the β-coefficient provides an estimate of the average yearly adjustment in r over the crisis episode.

Following the reasoning outlined in Roberts and Whited (2012), our empirical analysis rests on two identifying assumptions:

(A1) In the absence of the crisis, average price changes would have been the same across firms, irrespective of their degree of trade credit issuance.

(A2) There is no omitted variable correlated with trade credit issuance that affects prices during the crisis.

We assess the plausibility of these assumptions in the following ways. Firstly, we test for differences in pre-crisis trends in firm-product inflation rates between firms with low and high trade credit issuance, respectively. Secondly, we control for an additional set of potentially important confounding factors. For example, Chevalier and Scharfstein (1996) and Gilchrist et al. (2017) document that firms' liquidity positions underlie countercyclical mark-ups. To control for such mechanisms, we estimate augmented versions of Eq. (4) where we include interaction terms between the $Crisis$-variable and liquidity-related pre-crisis firm-characteristics. Finally, we test for cross-sectional heterogeneity in the effects of trade credit issuance on firm-product inflation rates with respect to firms' pre-crisis liquidity positions, and to increases in counterparty risk during the crisis, respectively. If the relationship between trade credit issuance and prices during the crisis period can be attributed to increases in liquidity costs and counterparty risk, then this relationship should be stronger for liquidity-constrained firms, as well as for firms that faced larger increases in counterparty risk.

Moreover, we assess the scope for an alternative explanation that increased trade credit demand per se lead to higher prices for firms that issued longer maturities during the crisis. This is done by estimating Eq. (4) using $\Delta Q_{i,p,t}$ as the dependent variable. In this specification, a rise in demand would be associated with a positive coefficient for $\hat{\tau}_i^{07}$, whereas an insignificant or negative coefficient would mitigate concerns for demand shifts underlying our baseline result.

4.2 The role of changes in trade credit contract maturities

A firm may, as noted above, respond to increases in liquidity costs and counterparty risks by raising the implicit interest rate in its trade credit lending, by reducing the maturity on this lending, or by some combination of the two. Any relationship between r and π^T will be attenuated if firms partly respond by lowering trade credit maturities, cf. Eq. (1). We do not observe contracted trade credit maturities in our data, so the results in our empirical analysis concern the effects of r on π^T, net of any changes in τ. This means, that the more firms experiencing increases in liquidity costs or counterparty risks in the crisis respond by reducing contracted maturities—rather than increasing the trade credit interest rate—the smaller will the coefficient β in Eq. (4) turn out. In the extreme, if all firms were to reduce maturities to zero—and effectively turn to cash transactions with trade credit premia set to zero—our estimated effect should be negative. The relative importance of increases in interest rates and decreases in maturities is ultimately an empirical question, and we stress that our analysis captures the net effect on prices of these two potential responses.

While we cannot disentangle the effects of changes in interest rates and maturities in our empirical analysis, we believe that there are grounds for assuming that trade credit contract maturities are relatively sticky—and that our results thus primarily capture the direct effect of r on π^T. We provide some evidence supporting this assumption in Figure 2, which shows average trade credit contract maturities, as measured by $\hat{\tau}_{i,t}$, within each tertile of the trade credit maturity distribution over the period 2004–2011. The figure in Panel A—in which we classify firms on a year-by-year basis, so that firms may switch between tertiles over time—shows that the the trade credit maturity distribution is stable over time; the average maturity within each tertile hardly changes at all between 2004 and 2011, with the exception of a very small decrease during the crisis.

In the figure in Panel B, firms are classified according to their position in the trade credit distribution in 2007, so that the set of firms in each tertile is fixed over time. This figure shows that trade credit maturities changed very little for the firms in the bottom and middle tertiles, but that the firms in the top tertile reduced maturities somewhat during the crisis, with $\hat{\tau}_{i,t}$ falling from 0.136 to 0.114 (i.e., from around 50 to 42 days). Thus, while there is scope for firms to change the maturity of their trade credit lending in response to shocks, the time-series presented in Figure 2 nonetheless suggest that trade credit contract maturities are overall sticky.

5 Results

Following the road map outlined in the previous section, we will begin our empirical account of findings by first presenting the baseline results, then move on to assessments of the plausibility of the two identifying assumptions, respectively, and finish with tests of cross-sectional heterogeneity in order to shed light on the underlying mechanisms.

As a starting point, Figure 3 provides an illustration of our main finding. It shows average firm-product inflation rates over the period 2004–2011 for firms with average trade credit maturities above (solid line) and below (dashed line) the sample median in year $t-1$. Inflation rates for the two groups of firms track each other closely in the four years leading up to the crisis, but then differ substantially during the 2008–09 crisis period. Although average inflation rates fall in both groups of firms—which is what one would expect in a crisis period with deflationary pressure—inflation rates fall considerably less among firms with long trade credit maturities. In the post-crisis period, inflation rates resume similarity across the two groups of firms. Thus, the figure provides initial support for our hypothesis that increases in liquidity costs and counterparty risk lead firms to raise trade credit premia. We will next substantiate by means of a formal analysis using the model specified in Eq. (4).

5.1 Main results

Table 2 reports the results for various estimations of the model specification in Eq. (4). The baseline result is reported in Column (I), showing a positive and statistically significant coefficient for the interaction term, $Crisis_t \cdot \hat{\tau}_i^{07}$. The coefficient indicates an average yearly increase in the trade credit interest rate, r, of 20.6 percentage points; resulting in an overall relative adjustment of 41.2 percentage points over the crisis episode.[16] In terms of implied price adjustments, comparing firms at the 75th and the 25th percentiles of the trade credit maturity distribution—a difference of around 23 days—amount to an average yearly difference of 1.3 percentage points and thus a two-year difference of 2.6 percentage points.

Next, we re-estimate the baseline specification using weights that adjust for differences in the shares of each firm's total sales accounted for by each of its products. More specifically, we estimate a weighted regression where the weight for each observation, $\omega_{i,p,t}$, is calculated as firm i's sales of product p divided by firm i's total sales. The results are reported in Column (II) and indicate a slightly smaller increase in the average yearly interest rate of 14.9 percentage points. Calculating price changes for firms at the 75th and 25th percentiles of the $\hat{\tau}_i^{07}$-distribution yields a relative price adjustment of 1.0 percentage points. These effects suggest that sellers more diversified product portfolios tend to make slightly larger price adjustments on average.

In spite of winsorization, there is still a concern for undue influence from a small number of firms with exceptionally long trade credit maturities. We therefore estimate a version of the baseline specification in which the main explanatory variable

[16]When replacing the time-invariant explanatory variable, $\hat{\tau}_i^{07}$, with lagged, time-varying maturities, $\hat{\tau}_{i,t-1}$, we obtain a coefficient of 0.188 (3.2) for the baseline specification.

is a dummy indicating whether a firm's trade credit maturity was above or below the sample median in the last pre-crisis year. The results, reported in Column (III), are consistent with the baseline results; the difference in inflation rates during the crisis between firms above and below the sample median of the trade credit maturity distribution is 1.3 percentage points. Similarly, one may be concerned that very large price adjustments drive the baseline result. To address the latter, we estimate a version of the baseline specification in which the dependent variable is replaced by a dummy that takes the value one for price increases, and zero otherwise. The estimated coefficient, reported in Column (IV), implies that firms with long trade credit maturities were 2.5 percentage points more likely to increase prices. These findings suggest that outliers in the dependent variable, or in the main explanatory variable, are not a concern for the baseline result.

The baseline specification includes firm-product-fixed effects to control for time-invariant differences in inflation rates across products. Hypothetically, time-varying differences in inflation rates across products could be important. Supposing that inflation rates during the crisis were higher for certain products, for reasons unrelated to trade credit issuance, and that the same products are customarily sold with long trade credit maturities, then our baseline result could be spurious. To address this possibility, we estimate a specification in which we replace the firm-product-fixed effects with firm- and product-year-fixed effects to control for the part of the variation in the inflation rate common to all producers of a given product. The resulting coefficient, reported in Column (V), is positive and statistically significant, with a magnitude of around half of the baseline coefficient. This seems to suggest that our baseline result is partly associated with time-varying product-specific factors. However, this saturated specification implies that remaining variation occurs in the sample segment involving two, or more, firms producing and selling a given product. Hence, the decline in effect magnitude—from 0.206 to 0.129—could well be due to an implicit change in sample composition, rather than the introduction of product-year-fixed effects. To test for this, we re-estimate the baseline specification on the sub-sample comprising two, or several, firms selling identically classified products. The obtained estimate is 0.162 (3.3); indicating that around half, 57 percent, of the decline can be attributed to a sample composition effect, and the other half to the introduction of product-year-fixed effects. If we instead estimate the baseline specification on the sub-sample of product-years involving a sole seller for given product, we obtain a coefficient of 0.399 (2.6); and the two estimates, 0.162 (3.3) and 0.399 (2.6), are statistically different at a ten-percent significance level. This difference in magnitude across the two sub-samples, suggests that the trade credit premia in part depends on competitive pressure—firms that face lower, as compared with higher, competition are more prone to adjust the trade credit price premium.

On a related note, we assess whether time-varying differences in inflation rates across industries could affect our results. We do this by estimating the baseline specification augmented with industry-year-fixed effects, where industries are defined using three-digit SNI/NACE codes. The result, reported in Column (VI), shows that the estimate of the coefficient β implies an average yearly increase in the trade credit interest rate of 17.5 percentage points.

Next, we tackle the possibility that our result could reflect a shift in demand during the crisis—away from sellers with short trade credit maturities and toward sellers with long maturities—as a result of longer trade credit maturities becoming more valuable for liquidity-constrained buyers during crises. To evaluate the demand-shift explanation, we regress the change in the quantity of sold goods, $\Delta Q_{i,p,t}$, on the right-hand side of the baseline specification. The idea is that an upward shift in demand for goods sold by firms with long trade credit maturities should cause an increase in both prices and quantities. Column (VII) shows, however, that the coefficient in this specification is negative and insignificant, which speaks against the alternative explanation based on shifts in demand.

Finally, we evaluate the parallel trends assumption by testing for differences in inflation rates between firms with long and short trade credit maturities in each year of the sample period. We do this using the baseline specification supplemented with interactions of the key explanatory variable, $\hat{\tau}_i^{07}$, and year-fixed effects. The resulting β-coefficients are plotted in Figure 4, using 2004 as base year. The coefficients are insignificant and close to zero in all pre- and post-crisis years, but positive and statistically significant in the two crisis years, which supports the plausibility of our first identifying assumption (A1).

Overall, the reported results are consistent with the conjecture that the surge in the cost of liquidity and counterparty risks during the crisis episode caused an upward shift in trade credit interest rates, and thereby product prices. The considerable magnitude of the interest rate effect could be viewed in light of opportunity costs of liquidity and default risk premia becoming potentially very large during the crisis episode (cf. Whited, 1992; and Berndt et al., 2018). Our estimate can further be related to the implicit interest rate on the "2/10 net 30" two-part terms contracts, which amounts to 45 percent (Giannetti, Burkart and Ellingsen, 2011). Thus, despite the large magnitude of our effect, it falls within the range of interest rates previously documented in the trade credit literature. We will next proceed to an evaluation of the plausibility of the second identifying assumption (A2).

5.2 The liquidity-constraint channel and other potentially confounding factors

Related to our second identifying assumption (A2), we now address the eventuality of firms' trade credit issuance being correlated with confounding factors that may give rise to a spurious relationship between trade credit maturities and price adjustments. We therefore re-estimate our baseline regression, while further controlling for firm-characteristics suggested by the literature to be potentially important determinants of firms' pricing behaviour. To this end, we augment Eq. (4) with interaction terms between the $Crisis$-variable and the additional factors under consideration—one at a time and jointly.

Extensive research has shown that liquidity constraints are a key determinant for firms' price setting. More specifically, Chevalier (1995), Chevalier and Scharfstein (1996), and Campello (2003) document that leverage leads to countercyclical price mark-ups.[17] To evaluate the scope for this channel in our context, Column (I) in Table 3 reports results when firms' pre-crisis leverage is accounted for. The coefficient

[17] Highly leveraged firms will be more vulnerable and exposed to liquidity shocks by way of a larger share of cash flow being committed to debt service.

for trade credit maturities remains positive and statistically significant; and for pre-crisis leverage we observe a positive and statistically significant coefficient, which is consistent with previous findings reported in the literature.[18] Another measure of liquidity constraints, used by Gilchrist et al. (2017), is cash holdings. Column (II) reports results from a regression where pre-crisis cash holdings is accounted for. The coefficients show that the effect for trade credit maturities persists, whereas the effect of cash holdings is statistically insignificant. However, consistent with the findings in Gilchrist et al. (2017), unreported regressions show that the effect of liquidity by means of cash holdings is negative and statistically significant in 2008. Thus, taken jointly, the results in Columns (I) and (II) suggest that the trade credit channel operates side-by-side with the direct impact of a mechanism induced by liquidity constraints.

Accounts receivables is a component of firms' working capital. A relevant question is therefore whether the channel we document is specific to trade credit provisioning, or whether inventories—the other main component of working capital—also played a similar role during the crisis. To examine this possibility we estimate a model where pre-crisis inventories is included. Column (III)-results show that the coefficient for trade credit maturities persists, but a negative and statistically significant coefficient for inventories. The latter result indicates that high-level inventory firms going into the crisis reduced prices, suggesting that they attempted to scale down their stocks.

Another factor that may influence our results is market power. That is, large firms may to a greater extent have increased prices during the crisis by means of dominant market positions. This may affect our trade credit maturity estimate if large firms tend to issue longer maturities. We account for market power by including the natural logarithm of firms' assets. The coefficients in Column (IV) for assets is negative and insignificant, whilst the trade credit maturity coefficient is largely unaffected.

Sales growth may also give rise to spuriousness for our baseline results. More

[18]Unreported results are very similar, if firms' short-term and long-term debts are included separately.

specifically, if sales are growing throughout the year, then the ratio of end-of-year accounts receivable to sales will lead to an overestimation of the true average payment terms; and vice versa if sales decline. We control for this by including pre-crisis sales growth. Column (V) shows that the inclusion of sales growth has little impact on the effect of trade credit maturities.

In the wake of the global financial crisis, world trade collapsed and Sweden, a small and open economy with exports around 50 percent of GDP, was severely hurt. This could potentially influence our results, since foreign trade typically is associated with long trade credit maturities, and exporting firms faced a severe drop in demand during the crisis. We therefore include a variable measuring the pre-crises export share of total sales at the 3-digit industry-level, $Export\ share_j^{07}$. Results reported in Column (VI) show that the coefficient for the export share is negative and statistically significant, whereas the trade credit maturities coefficient remains positive and statistically significant; and is even slightly enhanced in magnitude. These results may reflect that exporting firms experienced substantial drops in demand, imposing downward pressures on their prices.

In Column (VII), we report results when all factors are included jointly. The focal coefficient for trade credit maturities remains positive and statistically significant, while among the other six factors, leverage, inventories, and export share have statistically significant coefficients. Finally, Column (VIII) reports results where we further augment the model with pre-crisis measures of firms' tangible assets and profitability. Our estimate of the coefficient β in this specification is 0.169 (3.0), which is slightly smaller but still close to the estimate in the baseline specification.

Taken together, these complimentary findings indicate that the trade credit channel persists when we control for a wide set of potentially confounding factors. Moreover, leverage, inventories, and exports are in their own capacity important determinants of price adjustments during the crisis. The former yields an effect consistent with the framework outlined in Chevalier and Scharfstein (1996), stating that high leverage firms put a larger weight on current profits than future market shares in distressed periods. Thus, trade credit issuance makes for an additional channel through which financial frictions may influence prices—operating alongside the liquidity-constraint channel.

5.3 Cross-sectional heterogeneity

The finding that longer trade credit maturities cause larger increases in prices during the crisis should—according to the hypothesis outlined in the conceptual framework—be the result of some combination of increases in liquidity costs and in counterparty risks facing firms. To verify these mechanisms for our results, we conduct cross-sectional heterogeneity analyses in which we estimate the baseline specification on sub-samples of firms. The sample-splits are defined by empirical approximations of changes in liquidity costs and counterparty risk in the crisis. Our conjecture is that the association between trade credit issuance and price changes is stronger for firms that experienced larger increases in liquidity costs and counterparty risk, respectively.

We approximate for changes in liquidity costs using two measures of firms' pre-crisis liquidity positions: cash and liquid assets, $Cash/Assets^{07}_{i,t-1}$; and the size of unused credit lines, $Unused\ LC/Assets^{07}_{i,t-1}$.[19] Firms with weaker pre-crisis liquidity positions were presumably more vulnerable to the deterioration in access to external finance during the crisis, and can consequently be expected to have experienced larger increases in liquidity costs. Changes in counterparty risk are approximated using industry-level measures of changes in the average default probabilities in buyer industries during the crisis, $\Delta CP\ Risk^{07-09}_{j}$, where j denotes two-digit SNI/NACE industries.[20] For each variable, we construct two sub-samples: one with the firms in the bottom three deciles and one with the firms in the top three deciles.[21]

[19] We observe lending from the four major banks and therefore underestimate $Unused\ LC/Assets^{07}_{i,t-1}$ for firms obtaining credit from minor Swedish banks. This will lead us to underestimate the difference between the two sub-samples, if anything.

[20] The counterparty-risk measure is constructed as follows. First, we calculate the sales-weighted average default probability (PD) for each two-digit SNI/NACE industry and year. We then compute the change in each industry's weighted average PD between 2007 and 2009. Finally, we use the 2008-vintage of Statistics Sweden's input-output tables to calculate the industry-level measure of changes in average buyer PDs facing firms in each industry.

[21] The results are very similar if we instead define the sub-samples by splitting the sample at the median of each variable. The results from this alternative specification are available from the authors

The results from the cross-sectional heterogeneity analyses are reported in Table 4. Columns (I) and (II) cover the results for the sample splits based on the $Cash/Assets_{i,t-1}^{07}$-distribution. The coefficient is large and statistically significant for firms with low cash holdings, but relatively small and statistically insignificant for firms with high cash holdings; the difference is significant at the five-percent level. A similar pattern emerges in Columns (III) and (IV), where we report the results for sub-samples of firms with credit lines in the bottom and top of the $Unused\ LC/Assets_{i,t-1}^{07}$-distribution: the coefficient is large and significant in the former group, but smaller and insignificant in the latter; the difference is statistically significant at the ten-percent level in this case. The results reported in Columns (I)-(IV) thus support the notion that increases in liquidity costs account for some part of the positive relationship between trade credit issuance and price changes during the crisis.

The results for the sample splits based on changes in counterparty risk are reported in Columns (V) and (VI). The estimated coefficient is large and statistically significant in the sub-sample of firms that faced larger increases in counterparty risk during the crisis, but small and statistically insignificant in the group of firms for which the risk increase was smaller. The difference between the two coefficients is, moreover, statistically significant at the one-percent level, which suggests that increases in counterparty risk contribute to the positive relationship between trade credit issuance and price changes during the crisis. The results concerning counterparty risk should be interpreted with some caution, however. There are two main reasons for this. Firstly, $\Delta CP\ Risk_j^{07-09}$ is measured at the two-digit industry level and is by construction a crude approximation for the change in counterparty risk facing an individual firm. Secondly, $\Delta CP\ Risk_j^{07-09}$ is an ex post measure that in principle could be subject to reverse causality; this would be the case if price increases by sellers during the crisis—undertaken for reasons other than increased buyer credit risk—caused increased default risk for their buyers, rather than the other way around. These caveats notwithstanding, we believe that our results provide support for the hypothesis that part of our baseline results is accounted for by

upon request.

increased counterparty risk.

In sum, the results reported in Table 4 suggest that the positive relationship between trade credit issuance and price changes during the crisis is related to increases in liquidity costs and in counterparty risks. Thus, supporting that the baseline effect indeed is due to a rise in the trade credit interest rate, r.

6 Conclusions

Schwartz (1974) proposes the existence of a trade credit price premium, governed by an implicit interest rate determined by the selling firm's liquidity costs and the buying firm's default risk. This implies that increases in liquidity costs and counterparty risks should generate larger impacts on inflation rates for products sold by firms that extend more trade credit. By means of a difference-in-differences approach applied to Swedish manufacturing firm data, we relate adjustments in firm-product inflation rates in the 2008–09 recession to pre-crisis trade credit issuance. This approach enables a quantification of movements in implicit trade credit interest rates during the crisis period. We confirm that firms issuing trade credit with longer maturities exhibited substantially higher adjustments in firm-product inflation rates during the crisis; the baseline estimate indicate an overall rise in the annualized trade credit interest rate of 41.2 percentage points over the two-year crisis period. The documented effects are stronger for liquidity-constrained firms—which supports the notion that they reflect increases in firms' valuation of liquidity caused by contractions in the availability of external financing—as well as for firms whose buyers' creditworthiness shifted downwards. Hence, we find empirical support for the hypothesized determinants of the implicit interest rate in the trade credit price premium: liquidity costs and counterparty risks.

Our results relate to several strands of the literature. Firstly, the findings contribute to the trade credit literature by advancing the understanding of how trade credit is priced. More specifically, the notion that trade credit is an expensive form on short-term financing arises through the particularities of two-part terms contracts, in which buyers are offered discounts for early payments. For instance, in the well-known “2/10 net 30” contract, the buyer receives a 2 percent price discount for payments fulfilled within 10 days, and no discount for payments in the period 11 to 30 days—implying an annualized interest rate of 44.6 percent, if the buyer chooses to forego the discount and access credit in the 20-day payment period. Trade credit is nevertheless commonly issued under net terms contracting, for which no discount is offered, nor is an explicit interest rate specified. A valid question is therefore whether the widespread use of net terms contracts instead implies that trade credit is mostly an inexpensive funding source? According to our findings, the answer to this question is: not necessarily. Rather, sellers tend to compensate for the cost of extending credit by adding a premium to the product price. This premium may in turn become extensive in response to fluctuations in credit availability and counterparty risks.

Secondly, since the analysis concerns the role of financial frictions for the pricing of credit, our results further relate to the extensive literature on the bank lending channel. Previous work on bank lending suggest that quantity adjustments are of first-order importance during financial crisis episodes, whereas price adjustments tend to be less important. Our results suggest that price adjustments are of first-order importance for sellers that extend trade credit. These results can be reconciled in view of factors that favour quantity over price adjustments—such as adverse selection and moral hazard—being less important for trade credit issuance, as compared with bank lending. In other words, the stickiness of sellers customer bases in combination with goods being harder than cash to divert, make price adjustments more tractable.

Finally, our results further contribute to the growing literature on the influence of financial market imperfections on firms’ price setting. Notable work by Chevalier and Scharfstein (1996) and Gilchrist et al. (2017) show that liquidity constraints lead to countercyclical price mark-ups. This paper highlights that trade credit issuance induces an additional channel alongside the previously documented one—that can explain countercyclical movements in prices.

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Table 1: Descriptive statistics

This table reports descriptive statistics for all variables used in the empirical analyses, as well as for some additional firm characteristics. Definitions of the variables are provided in the text.

	Mean	Median	Std. dev	Pct. 10	Pct. 25	Pct. 75	Pct. 90	No. obs
Panel A. Price and quantity variables (2004–2011)								
Firm-product inflation ($\pi_{i,p,t}$)	0.028	0.006	0.161	–0.149	–0.032	0.092	0.224	49,134
Change in quantity sold ($\Delta Q_{i,p,t}$)	–0.008	0.005	0.456	–0.433	–0.150	0.156	0.395	49,134
Change in unit input costs ($\Delta UIC_{i,p,t}$)	0.037	0.022	0.350	–0.254	–0.075	0.137	0.322	49,134
Panel B. Key explanatory and sample split variables (2007)								
Trade credit maturity ($\hat{\tau}_i^{07}$)	0.097	0.094	0.058	0.022	0.061	0.125	0.162	3,928
$Cash/Assets_i^{07}$	0.086	0.024	0.130	0.000	0.002	0.116	0.272	3,928
LC_i^{07} (0/1)	0.604	1.000	0.489	0.000	0.000	1.000	1.000	3,928
$Unused\ LC/Assets_i^{07}$	0.043	0.003	0.071	0.000	0.000	0.063	0.136	3,928
Panel C. Other firm characteristics (2004–2011)								
Trade credit maturity ($\hat{\tau}_{i,t}$)	0.090	0.089	0.049	0.025	0.059	0.117	0.149	18,885
$Cash/Assets_{i,t}$	0.081	0.022	0.125	0.000	0.002	0.110	0.258	18,885
$Total\ debt/Assets_{i,t}$	0.127	0.030	0.165	0.000	0.000	0.233	0.388	18,885
$Tangible\ assets/Assets_{i,t}$	0.267	0.245	0.184	0.040	0.112	0.392	0.527	18,885
$Cash\ flow/Assets_{i,t}$	0.075	0.074	0.140	–0.071	0.013	0.146	0.234	18,885
$Inventories/Sales_{i,t}$	0.147	0.128	0.106	0.032	0.077	0.193	0.275	18,885
$\Delta Sales_{i,t}$	0.037	0.046	0.197	–0.181	–0.044	0.136	0.242	18,885
$Assets_{i,t}$ (in SEK 1,000)	282,522	57,417	796,530	14,022	25,702	159,289	539,728	18,885
$Sales_{i,t}$ (in SEK 1,000)	354,555	98,216	817,508	26,419	44,393	249,974	758,715	18,885

Table 2: Baseline results

This table reports results for estimations of various specifications based on Eq. (4). The dependent variable is the firm-product inflation rate, $\pi_{i,p,t}$, in all specifications except those in Column (IV), where the dependent variable is a dummy equal to one for price increases and zero otherwise, and in Column (VII), in which the dependent variable is the change in the quantity of goods sold, $\Delta Q_{i,p,t}$. The regression in Column (II) is estimated using WLS, where the weight for each observation, $\omega_{i,p,t}$, is calculated as firm i's sales of product p divided by firm i's total sales in year t. The product-fixed effects are based on 8/9-digit CN codes. The industry-fixed effects are based on 3-digit SNI/NACE codes. High vs. Low $\hat{\tau}_i^{07}$ is calculated as the difference in the dependent variable between firms at the 75th and 25th percentiles of the $\hat{\tau}_i^{07}$-distribution. The estimation period is 2004–2011 in all columns. t-statistics calculated using robust standard errors clustered at the firm-level are reported in parentheses. ***, **, and * denotes statistical significance at the 1, 5, and 10 percent levels, respectively.

	(I)	(II)	(III)	(IV)	(V)	(VI)	(VII)
Dependent variable:	$\pi_{i,p,t}$	$\pi_{i,p,t}$	$\pi_{i,p,t}$	$\pi^{+}_{i,p,t}$	$\pi_{i,p,t}$	$\pi_{i,p,t}$	$\Delta Q_{i,p,t}$
$Crisis_t \cdot \hat{\tau}_i^{07}$	0.206***	0.149***		0.402**	0.129**	0.175***	–0.117
	(3.6)	(3.0)		(2.5)	(2.5)	(3.3)	(–0.8)
$Crisis_t \cdot \hat{\tau}_i^{07,High}$			0.013**				
			(2.1)				
High vs. Low $\hat{\tau}_i^{07}$	0.013	0.010	0.013	0.025	0.008	0.011	–0.008
Year FE	Yes	Yes	Yes	Yes	No	Yes	Yes
Firm FE	No	No	No	No	Yes	No	No
Firm × Product FE	Yes	Yes	Yes	Yes	No	Yes	Yes
Product × Year FE	No	No	No	No	Yes	No	No
Industry × Year FE	No	No	No	No	No	Yes	No
Product and firm controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Weights	No	Yes	No	No	No	No	No
R^2	0.327	0.399	0.327	0.302	0.487	0.374	0.327
Number of firms	3,928	3,928	3,928	3,928	3,928	3,928	3,928
Number of observations	49,134	49,134	49,134	49,134	49,134	49,134	49,134

Table 3: The liquidity-constraint channel and confounding factors

This table reports results from an augmented version of Eq. (4), where interaction terms between the $Crisis$-variable and a set of additional factors are included. The dependent variable is the firm-product inflation rate, $\pi_{i,p,t}$, and the estimation period is 2004–2011, in all specifications. Additional factors indicates whether, or not, $Crisis \cdot Tangible\ assets/Assets_i^{07}$ and $Crisis \cdot Cash\ flow/Assets_i^{07}$ are included as control variables. All specifications include firm-product-fixed effects and year-fixed effects. t-statistics calculated using robust standard errors clustered at the firm-level are reported in parentheses. ***, **, and * denotes statistical significance at the 1, 5, and 10 percent levels, respectively.

	Dependent variable: $\pi_{i,p,t}$							
	(I)	(II)	(III)	(IV)	(V)	(VI)	(VII)	(VIII)
$Crisis_t \cdot \hat{\tau}_i^{07}$	0.190***	0.205***	0.226***	0.183***	0.193***	0.211***	0.193***	0.169***
	(3.4)	(3.5)	(3.9)	(3.3)	(3.2)	(3.8)	(3.5)	(3.0)
$Crisis_t \cdot Debt/Assets_i^{07}$	0.043**						0.034*	0.031
	(2.4)						(1.9)	(1.6)
$Crisis_t \cdot Cash/Assets_i^{07}$		–0.011					–0.028	–0.024
		(–0.5)					(–1.1)	(–1.0)
$Crisis_t \cdot Inventories/Sales_i^{07}$			–0.067**				–0.068***	–0.081***
			(–2.5)				(–2.6)	(–3.0)
$Crisis_t \cdot \ln(Assetsi_i^{07})$				–0.004			–0.002	–0.002
				(–1.4)			(–0.8)	(–0.9)
$Crisis_t \cdot \Delta Sales_i^{07}$					–0.013		–0.018	–0.012
					(–0.7)		(–0.9)	(–0.6)
$Crisis_t \cdot Export\ share_j^{07}$						–0.046***	–0.035**	–0.030**
						(–3.3)	(–2.5)	(–2.2)
Additional factors	No	No	No	No	No	No	No	Yes
R2	0.328	0.327	0.327	0.327	0.326	0.328	0.326	0.326
Number of firms	3,928	3,928	3,928	3,928	3,879	3,928	3,879	3,879
Number of observations	49,134	49,134	49,134	49,134	48,699	49,134	48,699	48,699

Table 4: Cross-sectional heterogeneity

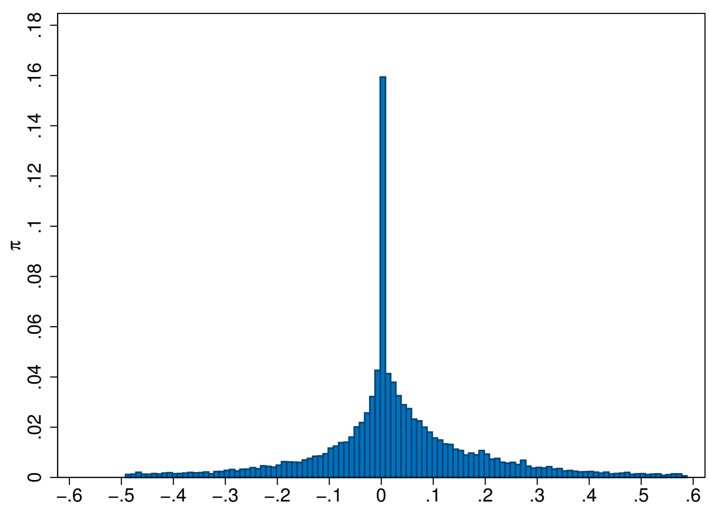
This table reports results for estimations of the baseline specification in Eq. (4) on various sub-samples of firms. Columns (I) and (II) report results for sample splits based on cash holdings; Columns (III) and (IV) for sample splits based on unused credit lines; and Columns (V) and (VI) for sample splits based on changes in counterparty risks. The cutoffs used to construct the sub-samples are defined at the firm level; hence, the number of firms in each sub-sample is approximately the same, while the number of observations differ somewhat. For the sample splits based on unused credit lines and changes in counterparty risk, the number of firms and observations also differ due to bunching of observations; in the former case because more than 30 percent of firms have no unused credit line, and in the latter because the cutoffs are defined using an industry-level measure. Reported p-values correspond to one-tailed tests, where the null hypothesis is that the β-parameters are equal in each pair, and the alternative hypothesis that the coefficients are larger in the groups of firms with low cash holdings, low credit lines, and high increases in counterparty risk, respectively. The estimation period is 2004–2011 in all columns. t-statistics calculated using robust standard errors clustered at the firm-level are reported in parentheses. ***, **, and * denotes statistical significance at the 1, 5, and 10 percent levels, respectively.

	Dependent variable: $\pi_{i,p,t}$					
	$Cash/Assets_i^{07}$		$Unused\ LC/Assets_i^{07}$		$\Delta CP\ Risk_j^{07-09}$	
	(I)	(II)	(III)	(IV)	(V)	(VI)
	Low	High	Low	High	Low	High
$Crisis_t \cdot \hat{\tau}_i^{07}$	0.367***	0.084	0.225***	0.082	0.062	0.453***
	(3.8)	(0.8)	(3.1)	(1.0)	(0.6)	(3.9)
p-value	0.021		0.092		0.008	
R^2	0.147	0.186	0.170	0.164	0.178	0.161
Number of firms	1,179	1,179	1,557	1,179	636	1,223
Number of observations	16,588	11,836	17,484	14,046	6,318	14,056

Figure 1: Firm-product inflation rates

Panel A of this figure shows the distribution of firm-product inflation rates in our sample over the entire sample period, 2004–2011. Panel B shows the average firm-product inflation rates in our sample, as well as the annual changes in the aggregate producer price index for the entire goods-producing economy (SNI/NACE sections A-E), for each year between 2004 and 2011. We calculate the latter as the log change in the annual average of the monthly values of the producer price index. *Source:* Statistics Sweden and authors' calculations.

A. Distribution of firm-product inflation rates in sample



B. Price changes in sample and in the aggregate economy

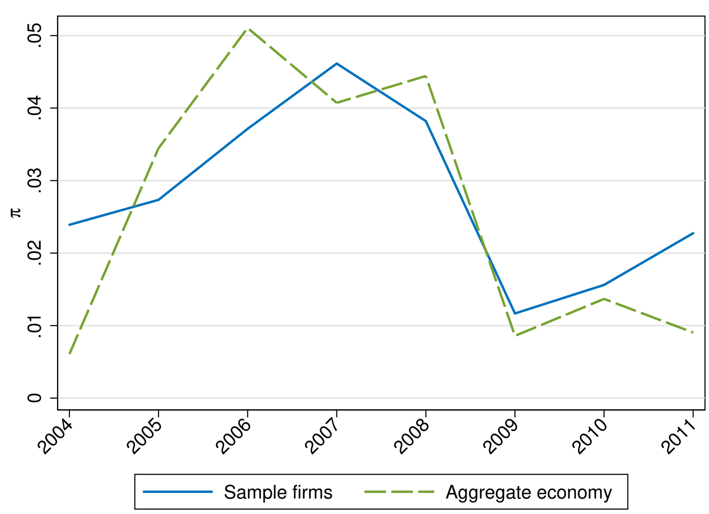
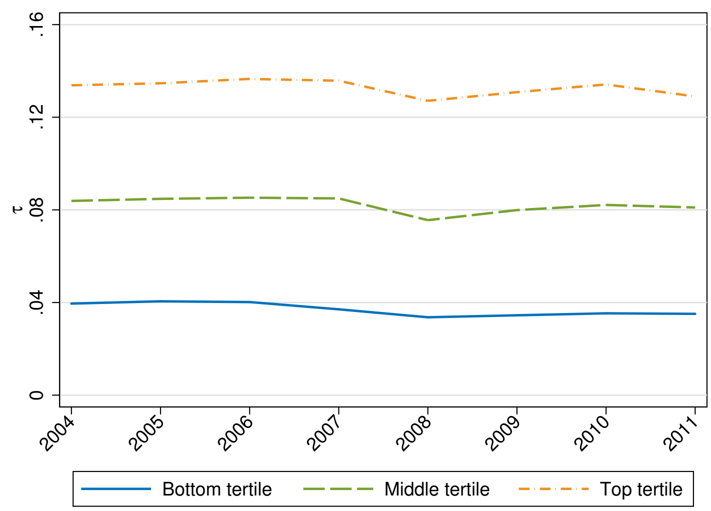


Figure 2: Average trade credit maturity by tertile

This figure shows average trade credit maturities, as measured by $\hat{\tau}_{i,t}$, within each tertile of the trade credit maturity distribution. The figure in Panel A shows firms classified on a year-by-year basis, while in the figure in Panel B firms are classified according to their position in the trade credit maturity distribution in 2007.

A. Year-by-year classification of firms into tertiles



B. Fixed classification of firms into tertiles

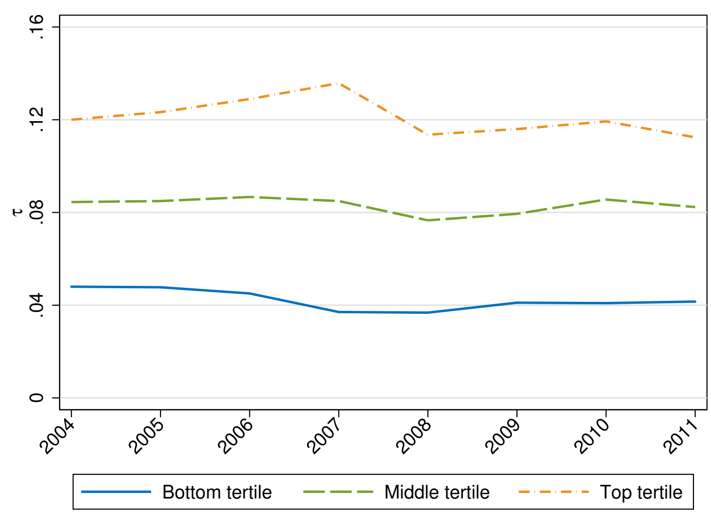


Figure 3: Average firm-product inflation rates over time

This figure shows average firm-product inflation rates in each year of the sample period for firms above (solid line) and below (dashed line) the median of the trade credit issuance distribution in year $t-1$, $\hat{\tau}_{i,t-1}$.

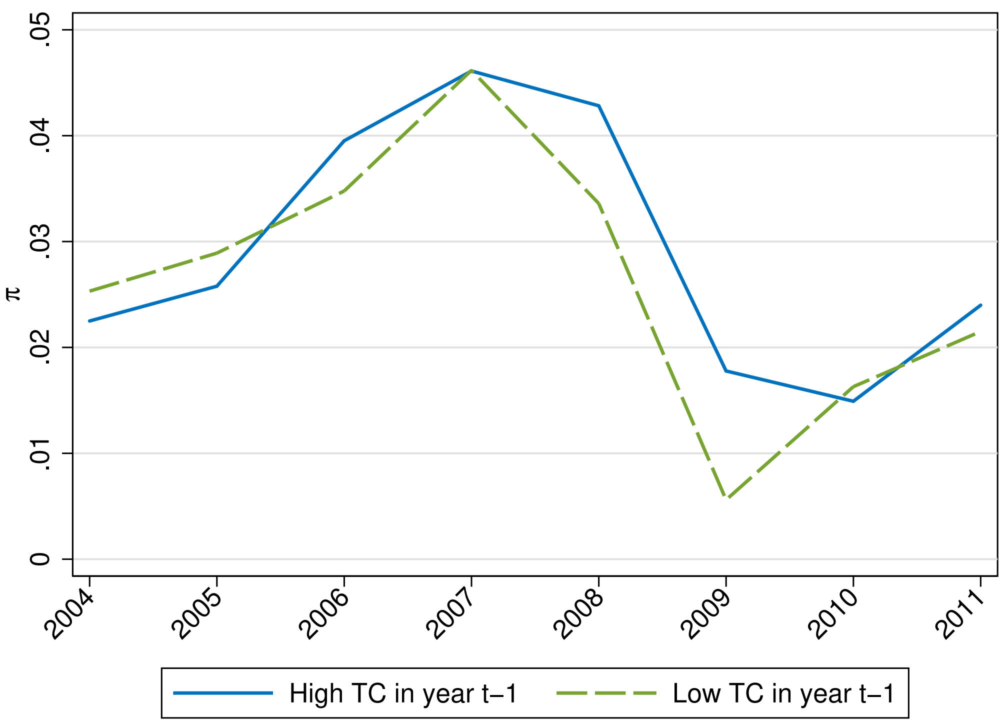
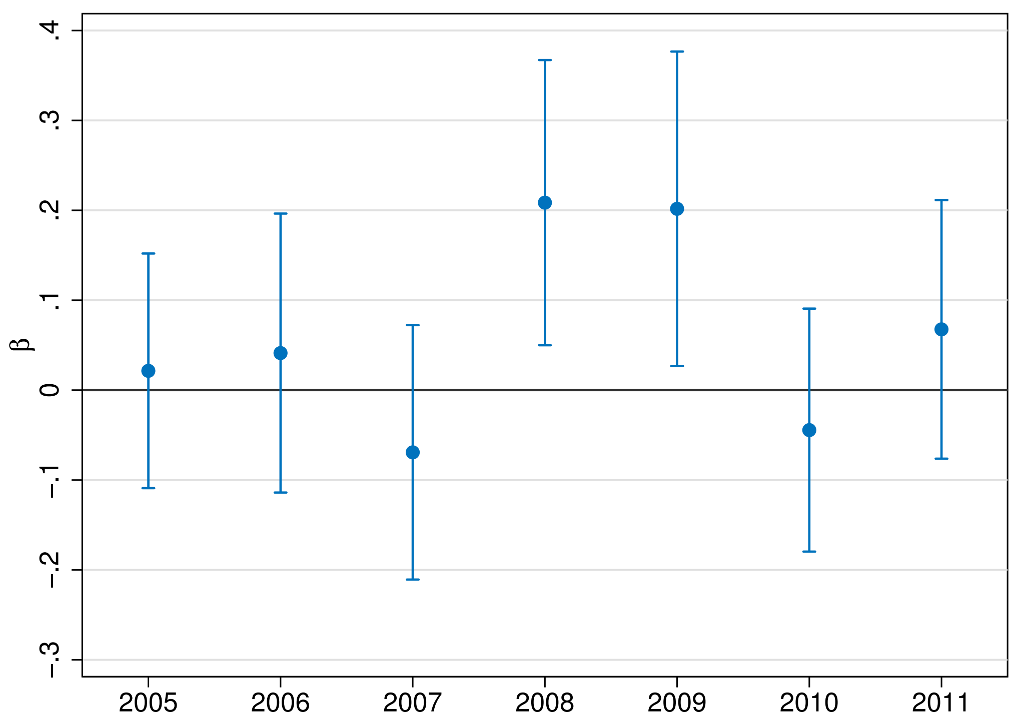


Figure 4: Pre-treatment trends

This figure shows the β_t-coefficients from an estimation of the baseline specification supplemented with interactions of the key explanatory variable, $\hat{\tau}_i^{07}$, and year fixed effects. The estimation is carried out using the entire sample period 2004–2011, with 2004 serving as base year. The estimating equation is thus: $\pi_{i,p,t} = \sum_{t=2005}^{2011} \beta_t \cdot Year_t \cdot \hat{\tau}_i^{07} + \alpha_{i,p} + \alpha_t + \gamma \cdot \mathbf{X_{i,p,t}} + \delta \cdot \mathbf{Z_{i,t-1}} + \varepsilon_{i,p,t}$. The vertical bars represent 95 percent confidence intervals.



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