Product Market Structure and Monetary Policy:
Evidence from the Euro Area[*]

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Abstract

Monetary policy aims at affecting corporate borrowing by influencing the marginal costs of firms, but its potency can be conditioned by the degree of market competition. We first identify conditions under which changes in marginal costs may have different effects on credit constraints and output under different competitive environment, in a simple Cournot competition setting. We then exploit changes in monetary policy to examine whether the pass-through of borrowing costs is affected by market structure. First, we use as an experiment the announcement of the ECB Outright Monetary Transactions (OMT) program in a triple-differences specification. We show that small firms (which have low market power and higher credit constraints) in "stressed" countries (which benefited more from the policy) within less concentrated sectors experienced a larger reduction in credit constraints than similar firms in more concentrated sectors. Second, we exploit continuous state-of-the-art measures of monetary policy shocks to study how market structure affects pass-through to real variables, like investment and sales growth. We find evidence that firms with more market power respond less to monetary policy shocks. These results show that the interaction of borrowing capacity and market structure matters, and that concentration may have important effects on monetary policy transmission.

1 Introduction

One of the most prominent academic and policy debates over the last few years has centered around the rise of market power. This can take the form of increasing concentration of sales to a small number of firms in each sector, or the ability of firms to systematically charge prices above marginal costs. Several studies have pointed to substantial growth along all these margins for the past two decades in the United States. This development has often been tied to a number of outcomes, such as stagnant investment, falling labor share, and increasing earnings inequality (Aghion et al. 2005, De Loecker et al. 2020, Gutiérrez and Philippon 2017, Díez et al. 2018, Syverson 2019, Barkai 2020). There is, though, much less agreement on the causes and consequences of these apparent trends (Autor et al. 2017, Traina 2018), while the issue seems to be less salient for Europe (Cavalleri et al., 2019).

Notwithstanding whether it is increasing or not, the very existence of some degree of market power among sectors in the economy may well have a bearing on how monetary policy impact activity. The implications of market power for monetary policy is one aspect of market power that has so far received very little attention in the literature (Scharfstein and Sunderam, 2017; Aghion et al., 2019; Duval et al., 2021), even though its existence has the potential to severely impair the transmission of monetary policy. Researchers have tended to see competition status as a given background effect that has no implication for the transmission of aggregate shocks and policy. However, more broadly taken, we have seen that the degree of competition can matter for the incentives to innovate and protect rents (e.g., Aghion et al. 2005, Gutiérrez and Philippon 2017, Vives 2008).

Standard theory posits that an immediate effect of monetary accommodation is that it reduces firms' borrowing costs, allowing them then to better manage their financial position such as to increase innovation, investment, and hiring, thereby boosting economic activity. This is the case for either the traditional money-view, where monetary policy operates through the expectations channel and affects the user cost of capital, or the credit-channel view, where monetary policy operates through its effects on the external finance premium, and hence corporate balance sheets, by an "accelerator" mechanism and an augmented cost of capital (Bernanke and Gertler, 1995). It is also the case for policy operating through conventional means, by manipulation of the short-term interest rate or through more recent unconventional policies, where central banks more directly attempted to influence long-term interest rates.

The potency, then, of monetary policy to affect aggregate demand depends on its ability to affect

firms' borrowing costs. As such, deviations from the perfectly competitive benchmark may influence this ability in a number of ways. First, note that firms produce at the point where marginal revenue equals their marginal cost; monetary policy can shift marginal cost and hence affect output through its effects on the cost of capital (over a sufficiently long period such that capital is variable). The pass-through of monetary policy is conditioned by the competitive conditions of the market. Indeed, the output response will depend on how marginal revenue (typically assumed to be decreasing in output) moves with market power. In particular, it may be presumed (e.g. Syverson, 2018) that in a perfectly competitive market, characterized by a flat marginal revenue curve, shifts of the marginal cost curve have a larger effect on output than under a case where firms have market power and face downward residual demand and marginal revenue curves.

This reasoning (however plausible and intuitive) needs to be qualified. If market power steepens the marginal revenue curve, then the output reaction to lower cost will be subdued. However, market power is identified with the slope of the residual demand faced by the firm, which is not the same as the steepness of marginal revenue. The impact of market power on pass-through depends on the *curvature* of the demand function. In particular, on whether the demand function is log-concave or log-convex. As such, the presence of market power under log-concave demand directly limits the pass-through of policy changes; a monetary contraction will lead to a lower reduction in output, and a monetary expansion will lead to a lower rise in output. Note that this simple analysis may, if anything in practice, underestimate the role of market power; some global firms have very high cash reserves and a concomitant borrowing capacity that small changes in interest rates may not in fact move their marginal cost curves at all.[1]

Second, the presence of market power by some firms may occur at the expense of other, "disadvantaged", or "squeezed", firms. Inter alia, firms with high market power may strike preferential agreements with upstream and downstream firms, and may face lower costs of advertising and retailing; market power may be the result of political connections, which could lead to reduced entry; and firms with market power in their product market may also yield such power in specialized labor markets, preventing their competitors from attracting skilled workers (Krueger and Ashenfelter, 2017).[2] All these imply that potential projects of disadvantaged firms may be less profitable than otherwise, resulting in reduced borrowing capacity relative to other markets. Put another way, the pass-through of monetary policy to lending conditions may be affected by market structure. This margin may be quantitatively crucial,

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[1] In other words, companies such as Apple or Google may face little to no external finance premium.

[2] The flip side, of course, is that firms may also exercise power in the market for their inputs; indeed, there is evidence of such a phenomenon for the United States (Azar et al., 2017).

given that small and medium-sized enterprises (SMEs) are responsible for a substantial fraction of economic activity across modern economies, and particularly in the euro area. Figure 1 shows the share of total value added in the euro area attributed to small and medium-sized firms. While the variation is substantial and the values are naturally higher in smaller countries, even in Germany and France the value added share of SMEs is 55%; it is 61% in Spain and 66% in Italy. Frictions to the transmission of monetary policy to SMEs can then severely curtail the degree to which policy can be an effective demand-management tool.

In this paper, we test both these hypotheses. After discussing in detail theoretical equilibrium conditions under which market power may hamper monetary policy pass-through (see below), we first empirically examine whether it can affect the transmission of monetary policy to *lending conditions*. We exploit variation in borrowing costs induced by the introduction of the European Central Bank's (ECB) Outright Monetary Transactions (OMT) program, formulated in August 2012. OMT was launched as a reaction to financial fragmentation across the euro area, with widely divergent borrowing costs for firms and households.[3] This divergence was considered to have severely impaired the transmission of ECB monetary policies across member states. The features of the OMT program which sought to address this fragmentation was no ex ante quantitative limits on secondary-market short-term bond purchases. The ECB had "pari passu" creditor status and there would be full sterilization of the liquidity effect. Moreover, applicable (or stressed) countries would be under a conditionality program.

The OMT's introduction effectively allows us to examine whether SMEs in less squeezed sectors experienced a large reduction in credit constraints. We use data from the Survey on the Access to Finance of Enterprises (SAFE), an ECB survey specifically designed to study access to finance for SMEs. The size and pervasiveness of the OMT shock provides a uniquely powerful experiment to test this hypothesis, given the relatively small sample size of SAFE.

It is important to note that by squeezed firms we do not mean firms that have weak profitability because they are low quality or have impaired balance sheets. In general, firms that lack profitable projects will indeed face stricter borrowing terms, as will firms that are highly indebted and cannot roll over their debt, an effect that is especially strong in crisis times (Duval et al., 2020). Instead, the focus here is on identifying whether squeezed firms in sectors with market dominance face different borrowing constraints than firms in more competitive sectors.

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[3] For example, the spreads of Spanish and Italian ten-year government bonds relative to Germany increased by 250 and 200 basis points respectively in July 2012 compared with one year earlier.

Having looked at the impact of OMT on credit constraints, we then examine the pass-through of monetary policy to the *real economy*. We use a large cross-country firm-level database with detailed data on firms for the euro area, and state-of-the-art measures of monetary policy and examine how the effects of monetary policy shocks are conditioned by firms' market power. In particular, we use Local Projection techniques (Jorda, 2005), as implemented by Cloyne et al. (2019), and examine whether firms with low markups (and hence flatter residual demand curves) respond more to monetary policy shocks in their output and investment decisions, compared to firms with high markups. We measure markups using the canonical approach of De Loecker and Warzynski (2012). We sharpen identification by exploiting the finding that young firms are more credit constrained and hence respond more to monetary policy and the business cycle in general (Cloyne et al., 2019; Fort et al., 2013), and compare low markup young firm and high markup old firms.

We identify monetary policy shocks using high-frequency changes in interest rates before and after monetary policy decisions by the ECB, and use the framework of Jarociński and Karadi (2020) on the database of Altavilla et al. (2019) to disentangle the policy shock from the purely informational aspect of the announcement. This approach, though pioneered in a time series context, has become increasingly popular in work with microdata (Cloyne et al., 2019; Durante et al., 2020), and ensures that the shock only reflects exogenous variation in monetary policy and not feedback effects from the real economy to policy.

A particular concern for identification is the endogeneity of market power and concentration. A well-known fallacy in the industrial organization (IO) literature is comparing market outcomes to concentration, since concentration is itself a market outcome, and not necessarily indicative of market power. High market power in the strict sense of being able to charge prices above marginal costs may indeed be related to high concentration in a classic Cournot setting (where the market share of a firm is equal to the product of its Lerner index and the elasticity of demand). However, the incentives to enter into a market depend on the intensity of competition ex post. More competition upon entry will reduce profits and entry. Furthermore, in models where firms are heterogeneous in productivity and sell differentiated goods, higher competition may imply higher profits and concentration; higher competition reduces the range of productivity draws that allow profitable operation and hence reduces entry (see Syverson 2019).

At the same time, even if structural market characteristics imply that competition leads to high concentration, this could still be important for monetary policy transmission. Indeed, one can imagine a situation where concentration is high due to high fixed costs; this means that marginal costs, which

monetary policy can affect, will be lower as a share of total costs, and hence less relevant a factor in the investment decisions of firms. This has in fact been pointed out as a particular characteristic of digital technologies (Korinek and Ng, 2018). In this context, monetary policy potency may be diminished as a result of high concentration.

We find strong evidence in support of our hypotheses for both sets of exercises. We first show that OMT had a much stronger impact in reducing borrowing constraints of SMEs in less concentrated sectors. At the same time, borrowing constraints had risen more before OMT for these sectors, consistent with the idea that concentrated sectors are in general less sensitive to changing financial conditions. We also find a similar behavior for sectors where the excess pricing power of top versus median firms is lower, though this is somewhat less strong than for concentration. Similarly, we then show that firms with higher pricing power (in the sense of higher De Loecker and Warzynski (2012) markups) respond less to monetary policy shocks. In particular, high markup firms (defined as those in the top quartile of the markup distribution) have a $0.5 - 1$ percentage point smaller contractions for both investment and sales following a one standard deviation increase in the shock.

Our second contribution is to more precisely spell out the conditions under which market structure may hamper monetary policy pass-through. With symmetric firms producing a homogeneous product at constant marginal cost, we show that if expansionary monetary policy reduces marginal costs and demand is log-concave (which is the more empirically plausible case), then in more concentrated markets the impact of monetary policy is lower. That is the case because pass-through (the response of marginal cost to price) is incomplete with log-concave demand; as such, when the intensity of competition rises, price becomes more sensitive to changes in marginal cost, and monetary policy gains more traction.

We then consider a model of heterogeneous firms to study how different firms respond to monetary policy. In a Cournot set-up with a homogeneous product, we have two sets of firms, which differ only in their marginal costs. Starting from an equilibrium, we show that a monetary policy easing which reduces marginal cost c does indeed increase profitability (and reduce credit constraints), but the impact of OMT on small firms (assumed to have higher marginal costs) is larger the lower the concentration in a Cournot model under certain conditions. In particular, and expressing marginal cost reductions as either ad-valorem or unit subsidies, this holds in general for constant elasticity demand with an ad-valorem subsidy in c or with linear demand and a unit subsidy in c. In "mixed" models with constant elasticity demand and unit subsidies or linear demand and ad-valorem subsidies, the outcomes are more complex.

The above effects occur even though the cost reduction is proportional for both types of firms in

the ad-valorem model, or of the same magnitude in the unit subsidy model. In the ad-valorem case the cost advantage of large firms is reduced and in the unit subsidy case it is not affected. In those circumstances, the effect of the cost reduction is to increase weakly the market share of small firms and reduce concentration.

This analysis implicitly assumes that credit constraints fall with higher profitability, and that the reduction in marginal costs is the same for all firms. A more realistic case is one where large firms are able to borrow in the capital markets: if easing has a larger impact through markets, which is exclusively used by larger firms and at the same time, banks have impaired balance sheets, further amplifying this difference, then the effect on profitability of small firms would be weaker. That is, in more concentrated markets we would have a lower impact of easing on SMEs' profitability if the cost reduction multiplier of large firms is high enough.

Relationship to the literature The heterogeneity of monetary policy transmission by firm characteristics has long been of interest to the literature, although more so recently. In an early contribution, Gertler and Gilchrist (1994) found that larger firms respond to a negative monetary policy shock by tapping credit markets and accumulating inventories, while small firms, who typically lack such access, cut down on production and draw-down inventories. Cloyne et al. (2019) focus on listed firms and show that investment and borrowing of older companies is largely unaffected by a tightening, while they decline considerably for younger companies (especially non-dividend payers), whose borrowing is more collateral-based. Jeenas (2018) shows that the disparity of firm responses across firms is large, and that liquidity and leverage are important predictors of such discrepancies. In particular, highly levered or illiquid firms contract investment more after a negative monetary policy shock, but the importance of leverage goes away when controlling for liquid assets (who remained significant). Manea (2020) embeds financial constraint heterogeneity into a standard New Keynesian model and documents a more nuanced role for monetary policy than previously considered. Similarly, in the large literature on the investment effects of changes in the corporate tax (which also affects the cost of capital), while most recent research tends to find small aggregate responses (Yagan, 2015), other papers focusing on heterogeneous responses have found larger effects, in particular through financial frictions (Mahon and Zwick, 2017). In a different vein, Saidi and Streitz (2021) document a link between credit concentration and markups.

Moreover, the elasticity of demand (and thus the mark-up) affects the slope of the (New-Keynesian) Phillips curve (Galí, 2015); in fact, under the specific market structure underlying the standard New

Keynesian model (i.e. monopolistic competition, Calvo pricing, no strategic interactions), the Phillips curve steepens with higher markups. This implies that the sacrifice ratio falls, the cost of high inflation lower, which may lead to a different optimal rule, allowing for more absorption through inflation than output gap (Aquilante et al, 2019).[4] Similarly, the level of the markup, in so far as it impacts investment growth would also affect r^*, the equilibrium interest rate.[5] In terms of market concentration, if the economy is characterized by firms that have such significant profits and cash reserves that they become interest insensitive, then the slope of the Investment-Spending (IS) curve attenuates, rendering monetary policy less potent in moving aggregate demand (Aquilante et al, 2019).

Scharfstein and Sunderam (2017) was, to our knowledge, the first to consider product market power and pass-through, for the US mortgage industry. They show that high concentration in mortgage lending reduces the sensitivity of mortgage rates and refinancing activity to the yields of mortgage-backed securities – a key pass-through mechanism of quantitative easing (QE) in the United States. Aghion et al. (2019) analyze a question similar to ours, with different data. They examine how sectoral and firm growth after a monetary policy shock is conditioned by the interaction of liquidity constraints and competition. They find that more highly leveraged sectors benefit more from easing, and the more in less concentrated sectors, or countries with low product market regulation. Their firm-level analysis uses Worldscope, which has a very small number of small firms, hence potentially missing out on a great deal of the variation.

Closest to us is a contemporaneous paper by Duval et al. (2021), who also study how the transmission of monetary policy to the real economy is conditioned by markups in a similar setting. They also find evidence consistent with the hypothesis that firms with high markups respond less to monetary policy shocks. The main difference is that the monetary policy shock they use is based on forecast errors; on the other hand, they use a broader set of countries, but with similar data. As such, we deem their contribution complementary to ours. Furthermore, our paper differs in that we are able to show effects for credit constraints as well.

More broadly, our paper contributes to our understanding of how market power may impact the aggregate economy. While market power, and in particular markups, has traditionally featured prominently in models, including monetary models, the focus has typically been on its cyclical behavior. The findings of De Loecker et al. (2020), who link a number of phenomena to rising markups, have ushered

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[4]This may not be the empirically relevant case, however, see Del Negro et al. 2020

[5] A persistently lower r^* raises the probability of hitting the lower bound constraint. See also Bauer and Rudebusch (2020).

in a wave of new research. Covarrubias et al. (2010) argue that, post-2000, increasing concentration is linked with muted investment and lower dynamism in the United States. Eggertsson et al. (2018) show that a reasonable calibration of the neoclassical growth model to account for rising markups, together with falling real interest rates, can explain, inter alia, a falling labor share, an increase in the pure profit rate, and, crucially, a muted response of investment to rising Tobin's Q. Caballero et al. (2017) link rising rents, together with other aggregate trends, to rising risk premia. Azar and Vives (2019, 2021a,b) argue that the evolution of concentration alone cannot explain the documented aggregate phenomena for the U.S., unless models are augmented to account for the evolution of ownership structure (increase in institutional ownership and common ownership).

Organization   Sections 2 presents conditions under which product market conditions may affect the pass-through of monetary policy in a Cournot setting, first under homogeneous firms, then under heterogeneity in production efficiency. Section 3 discusses the data used in our analysis. Thereafter, in section 4 we explain our strategy to causally establish a connection between market power and monetary transmission (captured by the extent to which credit constraints bind) by exploiting a monetary policy shock caused by the ECB's Outright Monetary Transactions (OMT) program of the ECB. We estimate a triple differences specification and the results are shown in Section 5. We then move to examining real outcomes, with Section 6 outlining our estimation strategy and Section 7 presents the results. Section 8 concludes. Additional material is in the appendices.

2   Theoretical Determinants of Pass-Through

In this section, we discuss the underlying theory behind changes in firms' costs and the subsequent pass through to economic outcomes. We do so in the context of a Cournot model (the standard setup) allowing for firm symmetry and heterogeneity and alternative specifications of demand. The full underlying analysis for the material presented here can be found in the appendix.

2.1   The Determinants of Pass-Through in a Cournot Market

Our goal is to assess how a cost reduction induced by a relaxing of monetary policy impacts market outcomes. It may be presumed that pass-through will be impaired by market power in product markets. A firm with market power will produce to the point where its marginal revenue (MR) equals marginal

cost (MC) so, say, for vertical shifts of a constant MC, the output response will depend on how MR (assumed to be decreasing in output) moves with market power.

If market power steepens the MR curve, then the output reaction to lower costs will be subdued. However, market power is identified with the slope of the residual demand faced by the firm, and it must be noted that this is not the same as the steepness of MR. The impact of market power on pass-through depends on the *curvature* of the demand function. In particular, on whether the demand function is log-concave or log-convex. Furthermore, we need an equilibrium analysis to ascertain how changes in a market power parameter affects pass-through.

In what follows, we shall show that the result that market power impairs pass-through holds for a log-concave inverse demand, but not for a log-convex one. We analyze the determinants of pass-through in a symmetric Cournot market parameterizing market power by the degree of concentration (or inverse of number of firms in our case). We focus on the response of prices to costs (pass-through) to keep the analysis simple and be consistent with the approach of the literature; incomplete pass-through indicates that firms do not fully adjust output to the change in costs.

Consider a market for a homogeneous product with inverse smooth downward sloping demand function $P(Q)$. Let the relative degree of concavity of P, given by $\delta \equiv QP''/P'$, be constant. This is the case, for instance, for linear and constant elasticity demand. For $P' < 0$, P is log-concave for $1 + \delta > 0$ and log-convex for $1 + \delta < 0$.

Log-concavity of demand is standard in Cournot models. If it holds, then the elasticity of demand is increasing in price (elasticity of inverse demand is increasing in total output), the game is of strategic substitutes, and equilibrium exists. Furthermore, with constant marginal costs the equilibrium is unique (and symmetric), see Vives (1999).

Note however that with constant elasticity demand η, we have log-convexity since then $1 + \delta = -\frac{1}{\eta} < 0$. We obtain the following results when the $N \geq 1$ firms in the market have constant and equal marginal costs c and compete in quantities. Letting p^* be the equilibrium price, then the pass through is given by:

$$\frac{\partial p^*}{\partial c} = \frac{1}{1 + \frac{1+\delta}{N}}$$

With *log-concave* demand we have that: $\frac{\partial p^*}{\partial c} < 1$ and margin ($p^* - c$) decreases with c. With *log-convex* demand, by contrast, $\frac{\partial p^*}{\partial c} > 1$ there is over-shifting (that is, p responds more than c) and the margin ($p^* - c$) increases with c.

Note that these results hold for monopoly ($N = 1$) and perfect competition ($N \to \infty$). In particular, with perfect competition, we have that $\frac{\partial p^*}{\partial c} = 1$. As the intensity of competition increases (N grows), p^* becomes more (less) sensitive to changes in c if $P(Q)$ is log-concave (log-convex): $\frac{\partial p^*}{\partial c \partial N} > 0$ $\left(\frac{\partial p^*}{\partial c \partial N} < 0 \right)$. This means in particular that if demand is log-convex, pass-through with monopoly is *higher* than with perfect competition. A lower c increases profits if $2 + \delta > 0$. The upshot is that if expansionary monetary policy reduces marginal costs and demand is log-concave, then in more concentrated markets the impact of monetary policy is lower. Log-concavity, which implies that the elasticity of demand falls with output, is indeed the more empirically plausible case. The log-concavity assumption is in fact consistent with a range of empirical results in the trade literature in particular, where it is referred to also as Marshall's Second Law of Demand.[6]

2.2 Marginal Cost Reduction in a Heterogeneous Cournot Market

Now let us consider a Cournot market with two types of firms: low-cost and high-cost ones. The former will be large firms and the latter small firms. We examine how the impact of a relaxation of monetary policy on the profitability of high-cost firms, depends on the level of concentration in the market. See the appendix for detailed proofs.

Firms produce a homogeneous product and face an inverse downward sloping demand function $P(Q)$. We consider two cases:

Case 1: Linear demand

$$P(Q) = a - bQ$$

where $a, b > 0$.

Case 2: Constant elasticity demand

$$P(Q) = aQ^{-\epsilon}$$

where $a > 0$ and $0 < \epsilon < 1$.

There are two classes of firms with the following characteristics:

1. N_1 (small) firms with constant marginal cost c_1

2. N_2 (large) firms with constant marginal cost c_2

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[6] See Melitz (2018) and Mayer et al. (2020) for a summary of the empirical evidence and further discussion.

3. If $c_2 < c_1$. The cost advantage of low-cost firms is $\Delta = c_1 - c_2 > 0$.

Quantity is the strategic variable: firms of type i produces q_i amount of the product, $i \in \{1, 2\}$. A market equilibrium is a Cournot-Nash equilibrium of the game with $N_1 + N_2$ of players, (say with $N_1 > N_2$). Total supply with type-symmetric strategies is given by $Q = N_1 q_1 + N_2 q_2$. In both examples we have a unique Cournot equilibrium with type-symmetric strategies. In equilibrium, low-cost firms produce more than high-cost firms, they have lower market share and markups, and the Herfindahl-Hirschman Index (HHI) (i.e., the sum of squared firm shares) is decreasing with respect to c_2 (since increasing c_2 increases the market share of small firms s_1 at the expense of the market share of large firms s_2).

Suppose that monetary policy reduces marginal costs of all firms either by a factor $1 - \alpha$ (where $0 < \alpha < 1$), ad-valorem case, or as a unit subsidy $t > 0$, unit subsidy case:

– *Ad-valorem case*: High-cost firms' marginal cost becomes $(1-\alpha)c_1$. Low-cost firms' marginal cost becomes $(1 - \alpha)c_2$. Then the cost advantage is decreasing in α.

– *Unit subsidy case*: The effective cost of firm i is $c_i - t$. Then the cost advantage is independent of t.

Claim: We take as given that monetary policy easing reduces marginal costs, and assume that more profitable firms face lower credit constraints. Then in a Cournot model, under certain standard conditions, easing raises profitability, and the impact on small firms is larger the lower the market concentration. In the context of the model, this would require that the profits of small firms in equilibrium π_1^* rise less after easing when the marginal costs of large firms c_2 are lower (since in that case the concentration is higher); formally, $\frac{\partial^2 \pi_1^*}{\partial \alpha \partial c_2} > 0$ for the ad-valorem case, and similar for the unit subsidy case.

In order to test the claim, we identify the profitability prospects of SMEs with their level of perceived credit constraints. If the claim is true then the impact of easing in relaxing credit constraints should be subdued in more concentrated markets.

The claim is true in the "natural" models. In the constant elasticity model with an ad-valorem reduction, an increase in α increases profitability, and $\frac{\partial^2 \pi_1^*}{\partial \alpha \partial c_2} > 0$; therefore $\frac{\partial \pi_1^*}{\partial \alpha}$ decreases with a lower c_2. In the linear model with a unit subsidy, an increase in t increases profitability and $\frac{\partial^2 \pi_1^*}{\partial t \partial c_2} > 0$ where an increase in t reduces costs. In both cases an increase in α or t increases profitability for high and low-cost firms.

In the "mixed" models, the statements have to be qualified. In the linear ad-valorem case an increase in α also increases profitability for high-cost firms, but $\frac{\partial^2 \pi_1^*}{\partial \alpha \partial c_2}$ may be negative, although it is positive

whenever N_2 is high enough (there are several effects of an increase in c_2 on the marginal profitability of increasing α on π_1 with a positive effect of increasing c_2 on q_1 dominating the depressing effect of c_2 on the marginal impact of α on q_1 when N_2 is large). Furthermore, increasing α may even decrease the profits of low-cost firms. In the constant elasticity unit subsidy case, an increase in t need not increase the profitability for high-cost firms if t is high enough and/or c_2 low enough. In this case, a uniform subsidy to all firms may hurt high costs firms. However, whenever the unit subsidy benefits high costs firms, then we have also that $\frac{\partial^2 \pi_1^*}{\partial t \partial c_2} > 0$. See the model appendix. In all models a higher c_2 implies a lower market concentration as the market share of large firms is smaller.

The above effects occur even though the cost reduction is proportional for both types of firms in the ad-valorem model, or of the same magnitude in the unit subsidy model. In the ad-valorem case the cost advantage of large firms is reduced and in the unit subsidy case it is not affected. In those circumstances, the effect of the cost reduction is to increase weakly the market share of small firms and reduce concentration.

The effect would be stronger with a more positive impact differential of easing on large firms. This is particularly true of OMT, or in general of unconventional policies operating through financial markets. If these policies have a larger impact through financial markets, which are exclusively used by larger firms, and, at the same time, banks have impaired balance sheets, further amplifying this difference, then the effect of OMT on SMEs' profitability would be weaker (and in this case the effect of OMT could be to increase market concentration).

This can be modeled by OMT reducing marginal costs of large firms by a factor $(1 - \alpha)(1 - \beta)$, $0 < \beta < 1$: $(1 - \alpha)(1 - \beta)c_2$ in the ad-valorem model. The cost advantage of large firms grows with β: $\Delta = (1 - \alpha)(c_1 - (1 - \beta)c_2) > 0$. In this case, with linear demand, for example, we have that $\frac{\partial^2 \pi_1^*}{\partial \alpha \partial c_2} > 0$ if β is large enough. That is, in more concentrated markets we would have a lower impact of OMT on SMEs' profitability if the cost reduction OMT multiplier of large firms β is sufficiently high.

2.2.1 Constant Elasticity Model

In the constant elasticity model we have that in the Cournot equilibrium, $\pi_i^* = a\epsilon q_i^{*2} Q^{*-(1+\epsilon)}$. For the ad-valorem case, we have that:

$$\frac{\partial \pi_i^*}{\partial \alpha} = \frac{a(1 - \epsilon)}{(1 - \alpha)} q_i^{*2} Q^{*-(\epsilon+1)} > 0.$$

It follows that $\frac{\partial^2 \pi_1^*}{\partial \alpha \partial c_2} > 0$ since q_1^* is increasing and Q^* decreasing in c_2. And in the unit subsidy case we have the cost derivatives:

$$\frac{\partial q_1^*}{\partial c_2} = \frac{N_2 \big(Q^* - (\epsilon + 1) q_1^* \big)}{p^* \epsilon (N_1 + N_2 - \epsilon)} > 0$$

$$\frac{\partial \pi_1^*}{\partial c_2} = \frac{N_2 q_1^* \big(2 - (\epsilon + 1) s_1 \big)}{N_1 + N_2 - \epsilon} > 0.$$

However, in this case π_1^* need not be increasing in t if the cost advantage is large enough and/or t high enough. The reason is that whilst increasing the subsidy always benefits a low-cost firm, it need not benefit a high-cost one because the induced decrease in price hurts the latter more. This happens particularly if the subsidy is large. When this happens, high costs firms are hurt. In fact, q_1^* can be decreasing in t. However, whenever an increase in t benefits high costs firms, then we have also that $\frac{\partial^2 \pi_1^*}{\partial t \partial c_2} > 0$.

2.2.2 Linear model

We have that in equilibrium $\pi_i^* = b q_i^{*2}$. In the unit subsidy case we have:

$$\frac{\partial \pi_1^*}{\partial t} = 2 b q_1^* \frac{\partial q_1^*}{\partial t} = \frac{2 q_1^*}{N_1 + N_2 + 1} > 0,$$

(note that q_1^* is increasing in t). It follows that $\frac{\partial^2 \pi_1^*}{\partial t \partial c_2} > 0$ since q_1^* is increasing in c_2. In the ad-valorem case,

$$\frac{\partial \pi_1^*}{\partial \alpha} = 2 b q_1^* \frac{\partial q_1^*}{\partial \alpha} = 2 q_1^* \frac{N_2 \Delta + c_1}{(N_1 + N_2 + 1)} > 0$$

since q_1^* is increasing in α (the derivative with respect to α is $\frac{N_2 \Delta + c_1}{b(N_1 + N_2 + 1)} > 0$). We have that the sign of $\frac{\partial^2 \pi_1^*}{\partial \alpha \partial c_2}$ is ambiguous in general because $\frac{\partial^2 q_1^*}{\partial \alpha \partial c_2} < 0$. However, we have $\frac{\partial^2 \pi_1^*}{\partial \alpha \partial c_2} > 0$ if N_2 is large enough. The cross-derivative would also be positive with β high enough, meaning that the impact of monetary policy on costs is sufficiently advantageous for low cost (relative to high cost) firms.

3 Data

Before we come to the empirical analyses, we explain our data sources and treatment. We rely on two main data sources for our analysis. For the credit constraints exercises, our primary source of data is the Survey on the access to finance of enterprises (SAFE), a survey conducted on behalf of the European Commission and the ECB designed to examine access to finance for firms in the euro area, specifically focusing on small and medium-sized non-financial firms. The sample of interviewed firms is randomly selected from the Dun and Bradstreet database, and it is stratified by firm-size class, economic activity and country. The firms' selection guarantees satisfactory representation at the country level. For more information on SAFE, see Ferrando et al. (2019).

We then combine SAFE with a proprietary version of BvD Orbis (which we dub, SAFE-Orbis) to supplement SAFE with a more detailed sectoral classification.[7] We complement SAFE with measures of first sectoral concentration and then market power from CompNet, at the 2-digit level.

For **concentration**, we employ two commonly used measures:

1. the Herfindahl-Hirschman (HHI) which is the sum of squares of the firm share of sales in the relevant market, and;

2. the share of total sectoral sales accruing to the top 10 firms in the relevant sector (C_{10}).

For **competition**, we make use of four sectoral markup measures:

1. the profit margin;

2. the markup measure popularized by De Loecker and Warzynski (2012).

Our third and final markup proxy is the conventional price-cost margin, measured assuming either

3. variable (PCM_V), or;

4. fixed capital (PCM_F).

The price-cost margin is given by revenues minus materials and labor cost, divided by revenues, for PCM_V and revenues minus materials, labor cost and estimated capital cost, divided by revenues, for PCM_F. The profit margin is defined as operating profit over turnover. In particular, for each markup measure, we use the log of the difference between the markup at the 90th percentile of the sectoral

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[7] This is a mere subset of the better known Orbis database, which we use for the second part of the paper. See below for details.

distribution and the median markup. We do this as interest here is on the squeezing of firms that do not have power. As such, the relevant margin is the differential power of firms at the top versus firms in the middle.

Our investigation focuses on a sample of SMEs (firms with less than 250 employees) belonging to the following eight European countries (Belgium, France, Finland, Germany, Italy, the Netherlands, Portugal, and Spain)[8] observed from the first survey round (second part of 2009) until round 17 (first part of 2017). The sample includes 27,745 firms for a total of 51,347 firm-time observations.

Table 1 reports descriptive statistics on the main variables of interest, for the sample of firms that applied for a bank loan. Credit constraints, our main dependent variable, is a dummy variable that indicates firms as credit constrained if they report that:

1. Their loan applications were rejected;

2. Only a limited amount of credit was granted;

3. They themselves rejected the loan offer because the borrowing costs were too high, or;

4. They did not apply for a loan for fear of rejection (i.e., discouraged borrowers).

The indicator is equal to 1 if at least one of the above conditions (1-4) is verified, and 0 otherwise. By that definition, 11% of SMEs in our sample are credit constrained. The firms in the sample are almost all independent companies (94%); mature (82% are older than 10 years) and with relatively small annual turnover.[9]

The match with SAFE-Orbis allows us to get detailed information on the sectoral composition of the sample from the CompNet database, in order to construct the different measures of market power and concentration as explained below.[10] For the exercises on real effects, we use the full version of Orbis, the largest commercially available cross-country database on firm-level balance sheet data. While Orbis is not explicitly designed for research and several steps need to be taken to make it usable for our purposes, it has by now become a standard source of data for such studies, and we follow commonly established practices in treating the data, in particular following Kalemli-Özcan et al. (2019) to obtain firm level datasets comparable across countries. We first drop firms reporting negative total assets,

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[8] The selection of those countries is driven by the availability of the information on the sectoral characteristics.

[9] In particular, 41% of firms have annual turnover up to 1 million euro; 27% more than 1 million up to 2 million euro and 20% more than 2 million euro to 10 million euro.

[10] For details see https://www.comp-net.org.

negative employment, misreported employment (greater than 2 million employees), negative sales or negative tangible fixed assets.

We take particular care in deflating nominal variables using appropriate deflators, at the finest possible level of aggregation, following the approach of Gal (2013). We use deflators from OECD-STAN database, augmented where necessary with EU KLEMS, using separate deflators for output, value added, intermediate inputs and investment, at the two-digit industry level. Whenever output deflators are not available, we deflate output variables (sales and operating revenue) with the value added deflator. Given our data requirements, we end up with a sample of 8 euro area countries: Austria, Belgium, Finland, France, Germany, Italy, Portugal, and Spain.[11] Table 2 reports basic summary statistics for our main variables of interest. Our final sample contains around 1.1 million firms and, overall, we have almost 6 millions observations. The average firm size in the sample is small in terms of number of employees, 36 over the sample period. There are however also several large companies in our sample, and the largest ones have more than 200,000 employees. The mean of the investment rate is 12% and the cash flow to total assets is approximately 7%, which is a relatively low figure, implying that our firms are not so efficient in generating cash for future business activity. This is also reflected in the average low working capital ratio, which is equal to 22% of total asset. Our firms tend to use financial leverage to finance their activity, but on average it does not represent more than 30% of total liabilities.

Markup estimation We estimate markups using the canonical De Loecker and Warzynski (2012) method, a model-free approach which has become increasingly popular due to its minimal assumptions and low data requirements. The estimation uses the control function approach to work around the standard problem of simultaneity when estimating production functions, as input choices are a function of productivity shocks. The assumption is that lagged inputs are uncorrelated with current period shocks, and are used as instruments in GMM estimation, to recover output elasticities. Markups are defined over variable inputs, and so the choice of a variable input is crucial. Given that many European economies have relatively rigid labor markets, employment may not be flexible in the short run. As such, we follow an approach common in the literature (see e.g. García-Perea et al. 2020 or De Ridder 2019) and use materials as the flexible variable, and hence a gross output Cobb-Douglas production function,[12] using revenue as the output measure.

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[11] Of the original 12 countries of the euro area, we do not include Ireland and Netherlands, which are well-known to have poor representation in Orbis, Luxembourg because of the large presence of financial multinationals in the country, and Greece because very few firms report intermediates or value added.

[12] A translog production function yields very similar results.

We estimate the production function for each two-digit industry separately, pooling across all countries, but including country and year dummies. As detailed in Andrews et al. (2019), this allows for structural differences across sectors, while ensuring estimates are comparable across time and space through uniform output elasticities. At the same time, it provides for a much larger estimation sample in industries with few active firms. For further details on the methodology, see De Loecker and Warzynski (2012).

4   Empirical Framework for Credit Constraints

We attempt to causally establish a connection between market power and monetary transmission by exploiting a monetary policy shock caused by the Outright Monetary Transactions (OMT) program of the ECB, announced by President Draghi on July 26, 2012, which substantially reduced borrowing costs for sovereigns, and hence firms, in the euro area.[13] In particular, while OMT reduced borrowing costs across the euro area, it had a much larger impact on borrowing costs for countries in the euro area periphery ("stressed"). We hence implement a differences-in-differences-in-differences setup (DDD); we compare outcomes before and after treatment, between stressed and non-stressed countries, and by the degree of the sectoral structure (or market power) pw measure.

The model we estimate on a panel of bi-annual dimension time $(2011H2 - 2013H2)$ is as follows:

$$
\begin{aligned}
CC_{isct} = {} & \beta_0 + \beta_1 pw_{sct} + \beta_2 pw_{sct} \times OMT_t \\
& + \beta_3 pw_{sct} \times Stressed + \beta_4 pw_{sct} \times Stressed \times OMT_t \\
& + \mathbf{X}_{isct}\gamma + \lambda_{sc} + \lambda_{ct} + \lambda_{st} + \epsilon_{isct}.
\end{aligned} \tag{1}
$$

where OMT is a binary variable that takes the value of 1 after the policy treatment, and 0 before. The $Stressed$ dummy is equal to 1 for firms in Italy, Spain and Portugal, and 0 otherwise. A little more than one third of SMEs in our sample are in stressed countries. The continuous variable pw measures either the degree of concentration or pricing power, depending on the specification considered; higher values indicate higher concentration or pricing power, and the measures are defined below. We include country and sector dummies (λ_{sc}) to control for unobservable time-invariant characteristics of particular sectors, both separately and interacted; country-time dummies to control for country-specific trends (λ_{ct}); and sector-time dummies (λ_{st}) to control for disturbances affecting specific sectors globally.

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[13] For more details on OMT, see Ferrando et al. (2019).

We are essentially interested in the coefficient of the triple interaction term, β_4, which is expected to be positive if higher market power or higher concentration squeeze firms with low power.[14] While credit constraints as a whole fell substantially, if our hypothesis is correct, then for disadvantaged firms in these sectors, the pass-through of monetary policy would be weaker than observationally equivalent firms in other sectors, in stressed versus non-stressed countries. We are agnostic about the sign of the other β coefficients.

Term $\mathbf{X}$ includes a vector of relevant firm characteristics (i.e., pre-treatment characteristics) by industry, sector, and country. These are size (which we relate to employees and turnover), age, and qualitative firm characteristics. Size and age control for differential borrowing capacity for small and young firms, which may be unrelated to the presence of powerful firms in their sectors. The qualitative firm characteristics included are dummies for whether the firm's capital and credit history improved over the past six months, to control for firm-specific factors that could have affected credit demand and creditworthiness of potential borrowers, but unrelated to credit supply.

The identifying assumption here, as is standard with difference-in-difference-in-difference designs, is common trends, namely that, absent the shock, credit constraints would have developed similarly in relative terms across sectors, on average, within non-stressed relative to stressed countries. Sector or country-sector fixed effects control for unobserved heterogeneity at the sectoral level that could possibly confound this model, such as, for instance, structural determinants of financing conditions unrelated to market structure. We cluster errors at the level where we expect the shocks to operate, and hence at the country-sector level, as is common in these designs.

As mentioned, our interest is in examining the effects of both market power and concentration on monetary policy transmission. We use the C_{10} and HHI for concentration; for market power, we use price cost margin (assuming either fixed or variable capital), De Loecker-Warzynski markups and the profit margin. For each market power metric, we use the log of the difference between the median markup and the markup at the 90th percentile of the sectoral distribution. See Section 3 for details.

We limit attention to firms with less than 250 employees. While it is possible that even large firms may be squeezed by powerful competitors, large firms typically have better access to credit sources. Restricting attention to relatively small firms provides the most transparent way of creating meaningfully comparable treatment and control groups.

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[14] With the advance of micro-level datasets, triple interactions (as here) have become increasingly common in the economics and finance literature — see for instance Antràs and Chor (2013) and Beck et al. (2005).

We focus on credit constraints for two reasons. First, monetary accommodation is transmitted to the real economy through, inter alia, a reduction of firm credit constraints, and hence a natural outcome of interest is whether a firm is credit constrained or not. Second, credit constraint measures react quickly to news shocks (as was the case following OMT) and hence allow for sharper identification, particularly since we use biannual measures of credit constraints, relative to investment, which typically reacts more sluggishly. Of course, the ultimate goal of relaxing credit constraints is to increase aggregate demand, particularly investment, and hence the change in investment is a secondary and legitimate outcome of interest.

5 Results for Credit Constraints

Figure 2 shows the evolution of our financing constraint indicators for high and low concentration sectors (using both the C_{10} and HHI definition). The figure shows the time dummies from a regression of financing constraints on time dummies, firm controls, and country-by-sector fixed effects, for firms with less than 250 employees. If the hypothesis is correct and indeed squeezed firms in highly concentrated sectors experience a slower reduction in financing constraints after a positive monetary policy shock (which makes monetary policy more accommodating), then the opposite should also be true; that is, a deterioration of financial conditions should have a larger effect on small firms in sectors with low concentration. For firms that are squeezed in the first place and have relatively constrained access to finance, worse aggregate financial conditions should affect their constraints less.

The pattern in the figure shows precisely this. Financing constraints rose sharply for firms in low concentration sectors early in 2011 when the sovereign crisis took off, and then fell after OMT, exhibiting some volatility for the initial semesters. By contrast, financing constraints for similar firms in more concentrated sectors exhibited much shallower movements, albeit in a similar direction. Following the easing of financial conditions after OMT, there is another tightening in early 2014, again mostly concentrated in concentrated sectors, but the constraints quickly subside. Interestingly, financing constraints do not subside immediately after OMT is announced, but there is a delayed effect until early 2013.

Overall, credit constraints tighten faster and more severely in the less constrained sectors once the adverse shock hits, and also fall faster when monetary policy becomes accommodating. Additional programs by the ECB, such as targeted long-term refinancing operations (TLTROs), negative interest rates, and the various components of the Asset Purchase Program (APP), further eased credit conditions

and contributed to the additional reduction in credit constraints faced by SMEs, and benefited less concentrated sectors more. Note that the focus here is on the *evolution* of the financing constraints in this case, not the level, as it is possible that firms in different sectors have a different baseline in what they consider to be high or low financing constraints, which is why we also include sector fixed effects.[15]

Given the graphical evidence, we set our treatment period as given by survey rounds 9 and 10 (corresponding to 2013H1 and 2013H2), and rounds 6-8 (corresponding to 2011H2-2012H2) as the pre-treatment period. This is meant for expositional convenience only, as having more than two treatment periods would complicate the interpretation of the coefficients, with no difference in the results. In fact removing round 8 does not at all affect the results, precisely because the effect was delayed.

Tables 3-4 show results for specification (1) for the various concentration and power measures. Table 3 starts with the examination of the role of market concentration using C_{10}. Columns 1 and 2 show the results for specification (1), the first with separate country and sector fixed effects, and the second with country-by-sector fixed effects, purging our results from unobserved time-invariant effects. In the latter case, the effective assumption is that each sector is separate across countries; this is a more flexible specification, which however absorbs more variation, allowing us to make inference from changes within country-sector cells across time.[16]

We see that the results for (1) are very similar regardless of the fixed effect specification. We find that the coefficient of interest is positive, as hypothesized, and highly significant. While, controlling for concentration, the introduction of OMT introduced a sharp reduction in financing constraints (as evidenced by the graphical evidence), in particular in stressed countries, this was mitigated for firms in sectors with a high concentration of sales at the top. For firms in sectors with above median values (75th percentile) of C_{10} the reduction in financing constraints thanks to OMT was 2.5pp smaller than in sectors at the median. Columns 4-5 of table 3 repeats the same analysis using HHI as a measure of concentration instead of C_{10}. Results are very similar, and if anything stronger, as the coefficient of interest is significant across specifications.[17]

As is standard with DDD strategies, the identifying assumption is that, in the absence of the treatment,

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[15] The matched sample of SAFE and Orbis is too small for the first two rounds of SAFE covering 2009, and so we miss the initial part of the crisis.

[16] Note that all continuous variables are winsorized at 1%. The only exception is the sectoral markup measure, which is winsorized at 1% from below and 10% from above, due to a very long right tail.

[17] Recent literature (e.g., Cloyne et al. 2019) has shown that young firms, even if they are large enough to be listed, respond more to monetary policy shocks because they tend to be relatively more credit-constrained. SAFE is well-designed to study small firms as it oversamples them, but there is a much smaller coverage of young firms, resulting in samples less than 1/3 of the size presented in Table 2. Nevertheless, results are very similar.

the credit constraints of firms would have evolved in a parallel fashion, which may not be true; in particular, it is possible that credit constraints for firms in low concentration sectors in stressed countries were already improving relative to the control group for reasons unrelated to credit supply (e.g., due to profitable opportunities). We exploit the fact that our data allows us to test for this assumption, and focus on rounds 6 and 7 (October 2011 to September 2012), just before OMT was announced. As such, we can test for common or parallel trends within the pre-treatment period, and perform a standard falsification placebo tests, where the pre-period is round 6 and the post-period is round 7.

Columns 3 and 6 reports the triple interaction coefficients for both C_{10} and HHI specifications, respectively; in both cases, the coefficient is small and not statistically significant (and for HHI it even has the wrong sign). These estimates suggest that no significant difference in trends in credit constraints existed across firms in the control and treatment groups, for the year before the OMT announcement.[18]

Tables 4 instead show the results of specification (1) using four different measures for pricing power in the sector: the price cost margin where capital is assumed fix (PCM_F), where it is assumed variable (PCM_V), the De Loecker-Warzynski markup, and the profit margin. In each case, we use the log of the difference between the 90th percentile and the median of the measure, to capture the pricing power of the top firms versus the median firm in the sector. We do this because technological differences across sectors may imply different median markups. The pricing power specifications yield less strong results. The coefficients of interest always have the correct sign, but are only significant for the markup and the profit margin (the latter only at the 10% level). The sample is substantially smaller, however, for these two measures, so results should be taken with a grain of salt.

Overall we do indeed find, as hypothesized, that the pass-through of OMT is stronger in less concentrated sectors, and especially in the stressed economies targeted by OMT, although also to some extent even in the other economies. The evidence for various estimates of pricing power is somewhat weaker.

6 Monetary Pass Through and Real Effects

Up to now, we have focused on the effects of market structure on the pass-through of monetary policy to credit availability. We now turn to the direct effects of market influence on monetary transmission, which is the ultimate goal of monetary policy. In this case, the outcome of interest is not credit availability, but

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[18]The recent literature has considered other failures of the parallel trends assumption, relating to staggered treatment timing or non-constant treatment effects. Our setup involves a unique treatment timing and has a relatively short time-frame, and so is not threatened by such violations. See Sun and Abraham (2020) and Borusyak et al. (2021) for discussions of the former problem, and for de Chaisemartin and D'Haultfoeuille (2020) for the latter.

rather measures of real economic activity. This is because theory predicts that market power directly impacts output.

As detailed in the model of Section 2 market power does imply lower pass-through relative to perfect competition if demand is log-concave but not if it is log-convex. We do not have rich enough data to discern the curvature of demand. In fact, sector- and firm-specific assumptions would be needed for this endeavor. While we are agnostic on this point, our interest is in estimating the reduced form impact of market power on monetary policy pass-through, not structural parameters of market structure. As such, if a sufficient numbers of industries were such that higher market power induces a higher response to monetary policy, then this would be a result against the notion that market power impairs, in equilibrium, monetary transmission.

The size of the dataset allows for a more general estimation strategy than the event-study approach of the credit constraints strategy. A classic challenge in studying the effects of monetary policy changes on real outcomes is that monetary policy itself responds to the business cycle. As Cloyne et al. (2019) note, an additional complication in a micro panel setting is the need to isolate interest rate changes from other shocks that affect interest rates (which in an aggregate time series setting would be handled through structural VAR techniques). We exploit recent advances in the literature of monetary policy measurement and directly examine how the effects of externally measured monetary policy shocks on real outcomes are affected by market power.

The monetary policy shocks are high-frequency responses of the 6 month EONIA swap rate measured around short intervals surrounding ECB monetary policy announcements, using the database of Altavilla et al. (2019). We follow the insight of Jarociński and Karadi (2020) to disentangle the monetary policy shock itself from the news it confers to the markets about the central bank's assessment of the economy. For instance, a contractionary monetary policy announcement moving both equity prices and interest rates in the same direction would mean markets recognize that the central bank expects the economy to overheat, and is hence not recognized as a shock. By contrast, a true surprise tightening would tend to raise interest rates and reduce equity prices. As such, we measure the shock ϵ_{mp} as the change in the EONIA rate before and after the announcement and press conference only for observations where the change in EONIA has the opposite sign from the change in a broad equity market index (STOXX). One issue here has to do with the scaling of the shock; given that the period covered is characterized both by historically low interest rates and a large role for non-standard measures, there is no clear correspondence to a simple policy rate. As such, we transform the shock into mean deviation form, where one standard

deviation is 9 basis points.

We employ the Local Projection approach of Jorda (2005), as used in Cloyne et al. (2019) and Durante et al. (2020). We further follow their approach (pioneered in an early version of Cloyne et al. 2019) and increase the granularity and support of our data by exploiting the fact that there is some variation across firms in the month on which they report their accounts. As such, even though our data come from annual accounts, we can create variation even within years by setting the month as our unit of time and the shock as the 12-month moving sum.

We focus on two real measures, investment and output. The latter because theory predicts a direct effect on output; the former because this is the traditional channel of transmission of changes in the cost of capital to output. A growing firm may increase its sales so much if it does not expand the scope of its operations through higher production capacity. We use the total net investment rate as our measure of investment, as this is what matters most for future productive capacity (Kalemli-Özcan, Laeven and Moreno, 2019). Our measure of output is total sales.

The specification is

$$\Delta_h X_{i,t+h} = \gamma_i^h + \sum_{g=1}^{G} \alpha_g^h [Z_{i,t-1} \in g] + \sum_{g=1}^{G} \beta_g^h [Z_{i,t-1} \in g] \epsilon_{mp,t} + \epsilon_{i,t+h} \tag{2}$$

where $\Delta X_{i,t+h}$ measures the changes in the outcome variable over h periods after the shock. $\Delta X_{i,t}$ measures the change from $t-12$ to t; as the shock is measured over the preceding 12 months, this allows for short-term reactions as well. X is either the change in the net investment rate or the log change in sales, while ϵ_{mp} is the policy shock, defined as above. Z refers to the defining characteristic over which we split the firms, split in groups (buckets); the primary interest of course is on markups, and so we split the sample in high and low markup groups. These are defined as the top and bottom quartile of the markup distribution. The coefficients of interest are β_g^h. We further include, as controls, two lags (12 and 24 months) of the change in cashflow, leverage and working capital; additionally for the investment specification, we include two lags of the change in sales, to control for differential growth opportunities across firms. We follow the convention in the literature and define the dependent variable in terms of changes from $t-1$. As such, the period 0 response in practice refers to responses for the 12 months following the shock, and not an instantaneous response.

This approach allows for a completely general non-parametric modeling and hence great flexibility in the specification. Indeed, to further exploit the granularity of our data and the flexibility of the

modelling framework, Z can be broken down further across a number of dimensions. In particular, recent evidence has shown that monetary policy is transmitted especially strongly through young firms (Cloyne et al., 2019; Durante et al., 2020), who, lacking a long-enough credit history, face especially strong credit frictions. As such, Z will also refer to splitting the sample in groups given by the intersection of markup and age, in particular young firms with low markups vs old firms with high markups. This split is expected to give the largest estimate, according to our hypothesis.

To keep things simple and avoid creating many groups, we focus on a sub-sample of only low markup young firms and high markup old firms. We follow the literature and assign firms to the young group if they were created less than fifteen years before the observation's year, and old otherwise.[19]

We follow the literature and cluster at the time and firm-level. However, we note that, as the shock is aggregate and random, we should expect that only cross-section correlation of errors is a concern, and not serial correlation; as such, clustering in the time dimension should be sufficient in controlling for off-diagonal correlation of errors. Indeed, in our estimates, clustering at the time-dimension only and then at time and country-industry or firm yields very similar errors. The randomness of the shock further implies that we add further controls for efficiency purposes only.

7 Results for Real Effects

We now present results for the real effects specification, in Tables 5-6. Recall that our hypothesis is that firms with high market power respond less to monetary policy changes. As a positive shock indicates an unexpected tightening, our hypothesis is that the β_g^h coefficients will be more negative for the low markup firms. For expository purposes, we omit the coefficients for the middle groups, but we note that the coefficients are always in between the low- and high-markup groups.

Table 5 shows results for investment, where we split the sample into markup bins. Panel A shows results for the baseline specification, with country-sector-year dummies, removing the average effect for firms in these buckets. This is important to control, for instance, for the possibility that some sectors may have high average markups but tend to respond less to monetary policy shocks for structural reasons. This way, we only compare firms within sectors, differentiated only by their markup levels. For both low- and high- markup firms, there is a pronounced V-shaped response to the shock, with the trough at $h = 1$, and the effect is significant for two years after the shock, though it remains negative even after 4 years.

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[19] For instance, it is not clear how old firms with low markups or young firms with median markups should be classified.

As such, the aggregate impact has standard qualitative characteristics. Of course, we are interested in the difference between the two coefficients, and the t statistic of coefficient equality (t-diff) is shown below the results. The coefficient difference is highly statistically significant at $h = 1$, and although it falls thereafter, the absolute magnitude is always larger for the low-markup group.

Panel B goes more granular and adds country-sector-month dummies instead, hence fully neutralizing any temporal variation. In Panel A, some within-year variation remains; the specification in Panel B allows us to take full advantage of the granularity of our data. The high-markup group is now the omitted category (as the within-month variation is removed) and the coefficient for the low-markup bin shows the difference between the response of low- and high-markup firms. We now see that there is a statistically significant difference also for the first year of the shock ($h = 0$), while the difference for $h = 1$ is quite similar to Panel A, and is very significant. Interestingly, there is also a response at the medium term ($h = 4$). As the monetary shock is standardized, we have that a one standard deviation increase in the shock measure leads to an approximately 0.59 percentage points larger response of low- relatively to high-markup firms.

Having confirmed our hypothesis, we can sharpen our identification by focusing on the subsample of firms which are expected to be on the extremes as regards their response to changes in monetary policy. In particular, we focus on low markup young firms and high markup old firms. The attractive feature of focusing on this split, other than the fact that the literature has recognized large differences in responses to monetary policy by firm age, is the fact that age is fully predetermined as the shock hits, hence allowing for clean identification. The results are shown in Panel C. As expected, they are stronger than previously, and there is now in fact a large difference between the two groups even at $h = 0$.

Table 6 shows results for the exercises with log sales growth. The structure of the table is similar, as we start with country-sector-year fixed effects by markup group, and we progressively saturate the estimation to include country-sector-month fixed effects, and then where we only compare low markup young firms with high markup old firms. Sales can respond much faster than investment, as they can more quickly reflect demand side responses from households. Firms can also draw down their inventories or increase production through an increase in their variable input use (quantity and intensity of labor and intermediates), without necessarily raising investment. As such, it is conceivable that the response of sales is faster than investment in the short run.

Indeed, this is what the results show. In panel A, which includes country-sector-year dummies, and hence allows us to also estimate the coefficient for the high markup group (and not just the difference

between the groups), we see that the response is stronger for both groups at $h = 0$ and persists as such throughout the horizon. However, as the high markup group also responds faster, the overall difference is somewhat smaller than for investment, on impact, but persists longer, as the differences between the two groups are significant even two years after the shock. Overall, depending on the specification, low markup firms have a $0.5 - 1$ percentage points larger contraction in sales following a one standard deviation increase in the shock.

8 Conclusions

In this paper, we assessed the interaction between monetary policy and the structural conditions of the real economy. Expansionary monetary policy seeks to ease credit constraints and reduce borrowing costs. By acting on a firm's marginal costs in this way it encourages higher employment and investment.

However, for an individual firm, the strength of that effect will be conditioned by its competitive environment. The overall effect is, though, by no means clear-cut. We demonstrated, in a simple Cournot setting, that the theoretical impact of market power on pass-through depends on the curvature of the demand function: market power under log-concave demand directly limits the pass-through of policy changes. Moreover, the presence of market power by some firms may occur at the expense of others.

To distinguish between these different channels of influence, we exploit exogenous variation in borrowing costs induced by the ECB's OMT program on the credit availability of firms. We show that small firms (which have low market power and higher credit constraints) in "stressed countries" within less concentrated sectors experienced a larger reduction in credit constraints than those in more concentrated sectors. We then study how market structure affects pass-through to real variables, using measures of monetary policy shocks, and find evidence that firms with more market power respond less to policy shocks.

These issues take on particular salience given the influential literature on the extent of market power in advanced economies in recent decades. The results of this paper point to an added benefit from improving the competitive landscape. This is particularly important at a time when monetary policy is constrained by the zero lower bound. At the same time, the theoretical framework points to potential diminishing returns from unconventional monetary policy measures, to the extent that they disproportionately benefit larger firms. The interaction of policy with structural characteristics of the real economy

should remain a high priority for research.

These results also have important implications in particular for the euro area, which is dominated by small firms. As such, the muted impact of monetary policy from market concentration for this group of firms could be particularly important, further highlighting the importance of EU competition policy, single market initiatives to enhance EU-wide competition, as well as structural reforms to reduce barriers to entry and the advantage of incumbents. The results of the present paper highlight that, in addition to the direct effect of enhanced competition and market efficiency, all of these initiatives would also contribute positively to a smoother transmission of monetary policy. This positive externality of competition policies for monetary policy has thus far not been considered, and could be very important.

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Figure 1: Value-added share of SMEs, by country
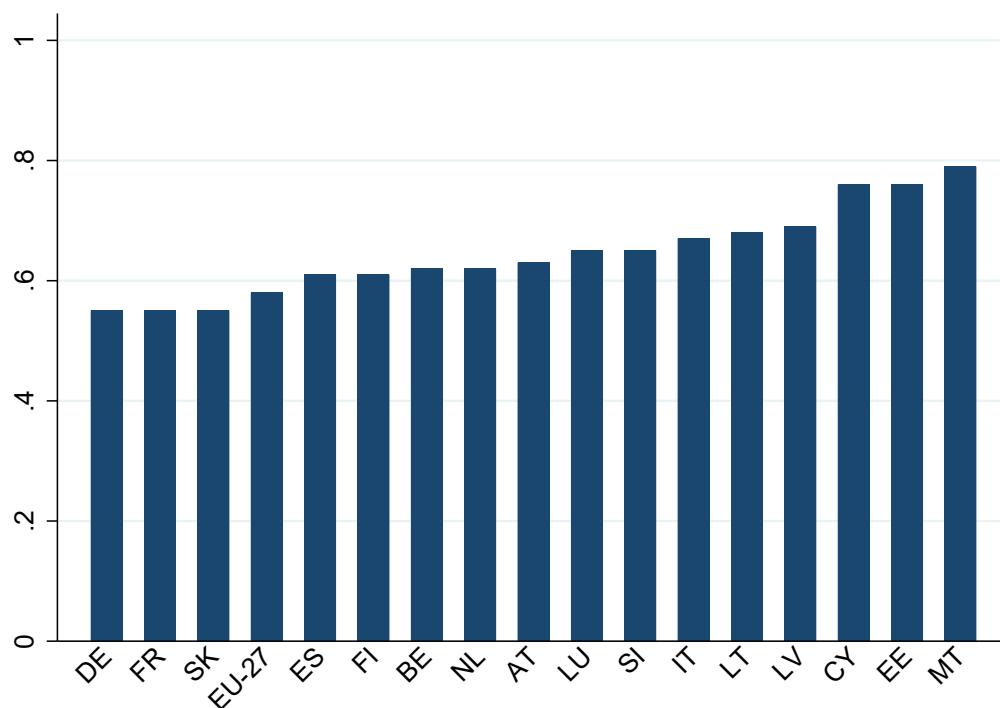
Figure 2: Evolution of financing constraints for high and low concentration sectors

(a) C_{10}

(b) HHI
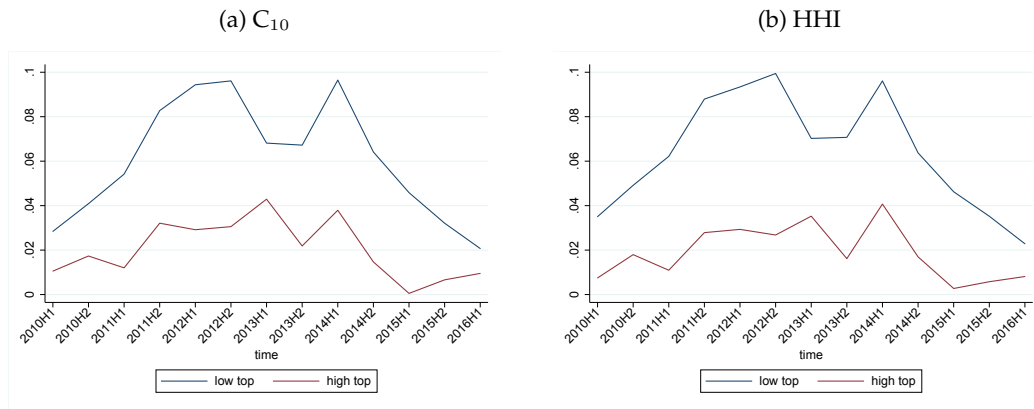
Notes: The figures show time dummy coefficients from regressing the credit constraint indicator on on time dummies, country-sector fixed effects, and firm characteristics, separately for firms in sector with high and low (above and below median) values of the respective concentration measures.

Table 1: Summary Statistics for SAFE sample

Variables	N	mean	sd	min	max
Dependent Variable					
Credit constraints	51,347	0.12	0.33	0	1
Firm Characteristics					
Size/type					
Autonomous	48,963	0.94	0.23	0	1
Micro	51,347	0.32	0.47	0	1
Small	51,347	0.36	0.48	0	1
Medium	51,347	0.32	0.47	0	1
Age (years)					
$\in [0,2)$	51,347	0.01	0.11	0	1
$\in [2,5)$	51,347	0.05	0.22	0	1
$\in [5,10)$	51,347	0.12	0.33	0	1
≥ 10	51,347	0.81	0.39	0	1
Stressed countries	51,347	0.42	0.49	0	1
Sectors					
Industry	51,347	0.30	0.46	0	1
Construction	51,347	0.12	0.32	0	1
Trade	51,347	0.26	0.44	0	1
Services	51,347	0.33	0.47	0	1
Financial Situation					
Turnover under 2m	51,347	0.45	0.50	0	1
Turnover 2-10m	51,347	0.30	0.46	0	1
Turnover 10-50m	51,347	0.20	0.40	0	1
Turnover 50m+	51,347	0.03	0.18	0	1
Improved capital	51,347	0.28	0.45	0	1
Improve credit history	51,347	0.27	0.44	0	1
Improve outlook	51,347	0.19	0.40	0	1
Market Structure					
C_{10}	39,856	0.21	0.18	0.012	0.90
HHI	42,179	0.02	0.03	0.0003	0.23
PCM_V	35,494	0.16	0.06	0.06	0.34
PCM_F	40,436	0.16	0.06	0.07	0.35
Markup	29,893	0.83	0.80	0.05	2.70
Profit Margin	26,345	0.08	0.04	0.00004	0.24

Notes: Source: SAFE Survey and CompNet. PCM_V, PCM_F, Markup and Profit Margin are defined as the (log) difference between the 90th percentile and the median in the sector for each measure. All market structure variables are winsorized at 1% top and bottom, except for Markup which is winsorised at 1% from below and 10% from above, due to a very long right tail.

Table 2: Summary Statistics for Orbis Sample

Variable	Obs	Mean	sd	min	max
Real Investment Rate	5,989,171	0.12	.69	-.69	6.36
Log Real Sales	5,876,101	14.37	1.52	0.05	19.26
Cashflow (% total assets)	5,989,171	0.07	.09	-1.34	1.39
Leverage (% total assets)	5,989,171	0.29	0.29	0	1
Working Capital (% total assets)	5,989,171	0.22	0.28	-0.99	0.99
DLW markups	5,989,171	1.63	0.87	0.73	5.63
Number of employees	5,989,171	36.04	575.45	2	257177

Notes: Source: BvD Orbis. All financial and income variables are winsorized at 1% top and bottom. Markups are trimmed top and bottom at 2%.

Table 3: Squeezed firms and concentration

Parameter/Measure	(1)	(2)	(3)[§]	(4)	(5)	(6)[§]
		C_{10}			HHI	
Measure	-0.0587	-0.3020		-0.0358	-0.4893	
	(0.0482)	(0.2031)		(0.1629)	(0.4331)	
Policy×Measure	0.1058	0.0644		-0.0020	-0.1576	
	(0.0662)	(0.0713)		(0.2569)	(0.2837)	
Stressed×Measure	-0.1243**	0.1289		-0.7945***	1.0917	
	(0.0548)	(0.4761)		(0.3053)	(1.8120)	
Policy×Stressed× Measure	0.2377***	0.2330***		1.5580***	1.8027***	
	(0.0734)	(0.0829)		(0.4763)	(0.5900)	
Placebo ×Stressed× Measure			0.0739			-0.1984
			(0.1270)			(-0.21)????
Controls	✓	✓	✓	✓	✓	✓
Country FE	✓			✓		
Sector FE	✓			✓		
Country×Sector FE		✓	✓		✓	✓
County×Time FE	✓	✓	✓	✓	✓	✓
Sector×Time FE	✓	✓	✓	✓	✓	✓
N	14298	14286	5519	15206	15189	5879
R^2	0.0609	0.0835	0.1091	0.0628	0.0860	0.1121

Notes: t statistics in parentheses, with standard errors clustered at the country-sector level. The significance stars are to be read as $* < 0.1,** < 0.05,*** < 0.01$. The dependent variable is the credit constraint measures. Each column shows results from least squares regressions of the dependent variables on the denoted controls, in addition to employment, age and turnover dummies, and dummies indicating whether the firm is a standalone entity, whether credit history improved over the previous wave, or whether the firm's capital situation has improved over the previous wave. [§] To avoid clutter, we simply report the interacted placebo coefficient in the Placebo case.

Table 4: Squeezed firms and pricing power

Parameter/Measure	(1) PCM_V	(2) PCM_F	(3) Markup	(4) Profit Margin
Measure	0.6181	-0.0885	-0.0291	1.1077
	(0.5625)	(0.3838)	(0.0732)	(0.7749)
OMT× Measure	-0.3750*	-0.4420**	-0.0187	-0.6711
	(0.2130)	(0.1906)	(0.0162)	(0.4760)
Stressed× Measure	-1.2701	0.3157	0.0476	-5.0500***
	(0.8290)	(0.7502)	(0.1379)	(1.8774)
OMT×Stressed×Measure	0.1868	0.3397	0.0376**	1.4385*
	(0.2281)	(0.2252)	(0.0171)	(0.7814)
Controls	✓	✓	✓	✓
Country×Sector FE	✓	✓	✓	✓
County×Time FE	✓	✓	✓	✓
Sector×Time FE	✓	✓	✓	✓
N	12809	14473	10910	8645
R^2	0.0854	0.0864	0.0868	0.1038

Notes: t statistics in parentheses. Standard errors clustered at the country-sector level. The significance stars are to be read as $* < 0.1, ** < 0.05, *** < 0.01$. The dependent variable is the credit constraint measures. Each column shows results from least squares regressions of the dependent variables on the denoted controls, in addition to employment, age and turnover dummies, and dummies indicating whether the firm is a standalone entity, whether credit history improved over the previous wave, or whether the firm's capital situation has improved over the previous wave.

Table 5: Investment - by markup group

	$h = 0$	$h = 1$	$h = 2$	$h = 3$	$h = 4$
A: Country-sector-year dummies, by markup group					
low markup	-0.1593	-1.1908***	0.0226	0.1099	-0.3902
	(0.3360)	(0.2853)	(0.1729)	(0.2145)	(0.2965)
high markup	0.1748	-0.6004**	-0.0415	0.0326	-0.0496
	(0.3647)	(0.2793)	(0.1887)	(0.1872)	(0.2894)
R^2	0.342	0.353	0.361	0.364	0.364
t-diff	1.88	3.68	0.26	0.29	1.47
N	5.9m	5.1m	4.3m	3.6 m	3.0m
B: Country-sector-month dummies, by markup group					
low markup	-0.4910***	-0.5880***	0.0464	-0.0066	-0.3922*
	(0.1837)	(0.1773)	(0.2541)	(0.2439)	(0.2349)
R^2	0.345	0.356	0.363	0.366	0.366
N	5.9m	5.1m	4.3m	3.6m	3.0m
C: Country-sector-month dummies, by markup by age group					
low markup & young	-2.6774***	-1.3847***	-0.6568	-0.6731	-1.1567**
	(0.4510)	(0.4908)	(0.8348)	(0.7678)	(0.5226)
R^2	0.340	0.353	0.364	0.368	0.369
N	1.4m	1.2m	1.0m	0.8m	0.7m

Notes: Standard errors in parentheses, clustered at the time and firm level. t-diff gives t statistics of difference in the indicated coefficients. The significance stars are to be read as $* < 0.1, ** < 0.05, *** < 0.01$. The dependent variable is the growth of the net investment rate, over $h + 1$ periods relative to t-1. Each column shows results from least squares Local Projections regressions of the dependent variables on the respective dummies, interacted with the shock, in addition to controls for sales growth, leverage growth, working capital growth and lagged dependent variable (all two lags), as well as two lags of the shock. The coefficients shown are the interactions of the low and high markup dummies with the shock (β_g^h from model (2)), omitting the middle group.

Table 6: Sales - by markup group

	$h = 0$	$h = 1$	$h = 2$	$h = 3$	$h = 4$
A: Country-sector-year dummies, by markup group					
low markup	-0.5610***	-1.7250***	-0.7346***	-0.0644	-0.4343
	(0.2056)	(0.4280)	(0.1971)	(0.3825)	(0.3936)
high markup	-0.3645**	-1.2149***	-0.3962**	0.1792	-0.2998
	(0.1650)	(0.3273)	(0.1810)	(0.2860)	(0.2871)
R^2	0.054	0.070	0.080	0.090	0.096
t-diff	2.11	2.56	1.40	0.89	0.52
N	5.9m	4.9m	4.2m	3.7m	2.9m
B: Country-sector-month dummies, by markup group					
low markup	-0.2387***	-0.4862**	-0.3900*	-0.2549	-0.1521
	(0.0791)	(0.2104)	(0.2129)	(0.2684)	(0.2729)
R^2	0.057	0.074	0.084	0.095	0.101
N	5.9m	4.9m	4.2m	3.7m	2.9m
C: Country-sector-month dummies, by markup by age group					
low markup & young	-0.3082***	-0.9530***	-0.8016***	-0.4345	-0.5371
	(0.0722)	(0.1793)	(0.2454)	(0.4150)	(0.4595)
R^2	0.066	0.087	0.099	0.111	0.121
N	1.4m	1.2m	1.0m	0.8m	0.7m

Notes: Standard errors in parentheses, clustered at the time and firm level. t-diff gives t statistics of difference in the indicated coefficients. The significance stars are to be read as $^* < 0.1, ^{**} < 0.05, ^{***} < 0.01$. The dependent variable is the growth of log sales, over $h + 1$ periods relative to t-1. Each column shows results from least squares Local Projections regressions of the dependent variables on the respective dummies, interacted with the shock, in addition to controls for leverage growth, working capital growth and lagged dependent variable (all two lags), as well as two lags of the shock. The coefficients shown are the interactions of the low and high markup dummies with the shock (β_g^h from model (2)), omitting the middle group; for Panel C the groups are low markup $\times$ young age and high markup $\times$ old age. Panel B includes country-sector-year dummies, and panel C country-sector-month dummies.

A Model Appendix

A.1 The Determinants of Pass-Through in a Cournot Model

Claim 1: *Pass-through (with constant marginal cost c). Let p^* be the Cournot equilibrium price and $\delta \equiv QP''/P'$ be constant. Then*

$$\frac{\partial p^*}{\partial c} = \frac{1}{1 + \frac{1+\delta}{N}}.$$

Proof. Immediate from first order condition when firms produce a positive output:

$$P(Q) + \frac{Q}{N}P'(Q) - c = 0.$$

Claim 2: *A lower c increases profits if $2 + \delta > 0$.*

Proof. Profits of firm i in equilibrium are given by

$$\pi_i^* = (p^* - c)\frac{Q^*}{N} = -\left(\frac{Q^*}{N}\right)^2 P'(Q^*).$$

It follows that

$$\frac{\partial \pi_i^*}{\partial c} = \frac{-(2+\delta)Q^*}{N(N+1+\delta)}.$$

Thus, we have that $\frac{\partial \pi_i^*}{\partial c} < 0$ if and only if $2 + \delta > 0$.

A.2 Heterogeneous Marginal Cost Reduction in a Cournot Market

Proposition 1.a *In the constant elasticity ad valorem model, we have that $\pi_i^* = a\epsilon(1-\alpha)^{-\frac{\epsilon+1}{\epsilon}}q_i^{*2}Q^{*-(\epsilon+1)}$*

$$\frac{\partial \pi_i}{\partial \alpha} = \frac{1-\epsilon}{\epsilon(1-\alpha)}\pi_i^* > 0,$$

and

$$\frac{\partial^2 \pi_1}{\partial\alpha\partial c_2} = \frac{1-\epsilon}{\epsilon}\frac{N_2}{N_1+N_2-\epsilon}\left(2 - (\epsilon+1)\epsilon(1-\alpha)^{\frac{\epsilon+1}{\epsilon}}s_1^*\right)q_1^* > 0$$

where $s_1^* = \frac{q_1^*}{Q^*}$.

Proof. From the FOC we obtain

$$q_1^* = (1-\alpha)\frac{-(N_2-\epsilon)c_1 + N_2c_2}{a\epsilon(N_1+N_2-\epsilon)}Q^{*\epsilon+1}$$

$$q_2^* = (1-\alpha)\frac{N_1c_1 - (N_1-\epsilon)c_2}{a\epsilon(N_1+N_2-\epsilon)}Q^{*\epsilon+1}$$

and the total quantity produced in equilibrium

$$Q^* = N_1q_1^* + N_2q_2^* = (1-\alpha)^{-\frac{1}{\epsilon}}\left(\frac{N_1c_1 + N_2c_2}{a(N_1+N_2-\epsilon)}\right)^{-\frac{1}{\epsilon}}.$$

Substituting this back into optimal quantities, we obtain

$$q_1^* = (1-\alpha)^{-\frac{1}{\epsilon}} \frac{-(N_2 - \epsilon)c_1 + N_2 c_2}{a\epsilon(N_1 + N_2 - \epsilon)} \left(\frac{N_1 c_1 + N_2 c_2}{a(N_1 + N_2 - \epsilon)} \right)^{-\frac{\epsilon+1}{\epsilon}}$$

$$q_2^* = (1-\alpha)^{-\frac{1}{\epsilon}} \frac{N_1 c_1 - (N_1 - \epsilon)c_2}{a\epsilon(N_1 + N_2 - \epsilon)} \left(\frac{N_1 c_1 + N_2 c_2}{a(N_1 + N_2 - \epsilon)} \right)^{-\frac{\epsilon+1}{\epsilon}}.$$

The price in equilibrium will be

$$p(Q^*) = aQ^{*-\epsilon} = (1-\alpha)\frac{N_1 c_1 + N_2 c_2}{N_1 + N_2 - \epsilon}.$$

It follows that

$$\pi_i^* = a\epsilon(1-\alpha)^{-\frac{\epsilon+1}{\epsilon}} q_i^{*2} Q^{*-(\epsilon+1)},$$

$$\frac{\partial \pi_i^*}{\partial \alpha} = \frac{1-\epsilon}{\epsilon(1-\alpha)}\pi_i^* > 0 \quad (\text{since } 0 < \alpha < 1,\ 0 < \epsilon < 1),$$

$$\frac{\partial^2 \pi_1^*}{\partial \alpha \partial c_2} = \frac{1-\epsilon}{\epsilon(1-\alpha)} \frac{\partial \pi_1}{\partial c_2} = \frac{1-\epsilon}{\epsilon} \frac{N_2}{N_1 + N_2 - \epsilon} \left(2 - (\epsilon+1)\epsilon(1-\alpha)^{\frac{\epsilon+1}{\epsilon}} s_1^* \right) q_1^* > 0.$$

Proposition 1.b. *In the constant elasticity demand with unit subsidy model, we have that $\pi_i^* = a\epsilon q_i^{*2} Q^{*-(\epsilon+1)}$. Both the equilibrium quantity and profit of the high-cost firm are increasing in c_2:*

$$\frac{\partial q_1^*}{\partial c_2} = \frac{N_2 \left(Q^* - (\epsilon+1)q_1^* \right)}{p^*\epsilon(N_1 + N_2 - \epsilon)} > 0 \text{ and } \frac{\partial \pi_1^*}{\partial c_2} = \frac{N_2 q_1^* \left(2 - (\epsilon+1)s_1^* \right)}{N_1 + N_2 - \epsilon} > 0.$$

An increase in t increases the profitability of the high-cost firm whenever t is low enough:

$$\frac{\partial \pi_1^*}{\partial t} = \frac{\left((N_1 + N_2)(\epsilon+1)s_1^* - 2\epsilon \right) q_1^*}{N_1 + N_2 - \epsilon} > 0 \iff t < \bar{t} \equiv c_2 - \frac{\left((\epsilon+1)(N_2 - \epsilon) + 2\epsilon^2 N_1 \right)\Delta}{\epsilon(1-\epsilon)}.$$

We have that

$$\frac{\partial^2 \pi_1^*}{\partial t \partial c_2} \propto \left(2 - (\epsilon+1)s_1^* - \frac{(\epsilon+1)(N_1 + N_2)s_1^{*2} - 2\epsilon}{(\epsilon+1)(N_1 + N_2)s_1^* - 2\epsilon} \right) \frac{\partial \pi_1^*}{\partial t},$$

which is positive whenever $\frac{\partial \pi_1^*}{\partial t} > 0$, and that

$$\frac{\partial \pi_2^*}{\partial t} = \frac{\left((N_1 + N_2)(\epsilon+1)s_2^* - 2\epsilon \right) q_2^*}{N_1 + N_2 - \epsilon} > 0.$$

Proof. We have that the equilibrium quantities fulfil

$$q_1^* = \frac{\epsilon(c_1 - t) - N_2\Delta}{a\epsilon(N_1 + N_2 - \epsilon)} Q^{*\epsilon+1}$$

$$q_2^* = \frac{\epsilon(c_2 - t) + N_1\Delta}{a\epsilon(N_1 + N_2 - \epsilon)} Q^{*\epsilon+1}$$

where $\Delta \equiv c_1 - c_2$, with the total quantity produced in equilibrium

$$Q^* = N_1 q_1^* + N_2 q_2^* = \left(\frac{N_1(c_1 - t) + N_2(c_2 - t)}{a(N_1 + N_2 - \epsilon)} \right)^{-\frac{1}{\epsilon}}.$$

It follows that

$$q_1^* = \frac{\epsilon(c_1 - t) - N_2 \Delta}{a\epsilon(N_1 + N_2 - \epsilon)} \left(\frac{N_1(c_1 - t) + N_2(c_2 - t)}{a(N_1 + N_2 - \epsilon)} \right)^{-\frac{\epsilon+1}{\epsilon}},$$

$$q_2^* = \frac{\epsilon(c_2 - t) + N_1 \Delta}{a\epsilon(N_1 + N_2 - \epsilon)} \left(\frac{N_1(c_1 - t) + N_2(c_2 - t)}{a(N_1 + N_2 - \epsilon)} \right)^{-\frac{\epsilon+1}{\epsilon}}.$$

The price in equilibrium will be

$$p(Q^*) = aQ^{*-\epsilon} = \frac{N_1(c_1 - t) + N_2(c_2 - t)}{N_1 + N_2 - \epsilon}.$$

The equilibrium profit of firm i is given by

$$\pi_i^* = a\epsilon q_i^{*2} Q^{*-(\epsilon+1)}.$$

We find that:

$$\frac{\partial q_1^*}{\partial c_2} = \frac{N_2 \left(Q^* - (\epsilon + 1)q_1^* \right)}{p^* \epsilon (N_1 + N_2 - \epsilon)} > 0,$$

$$\frac{\partial \pi_1^*}{\partial c_2} = \frac{N_2 q_1^* \left(2 - (\epsilon + 1)s_1^* \right)}{N_1 + N_2 - \epsilon} > 0,$$

$$\frac{\partial \pi_1^*}{\partial t} = \frac{\left((N_1 + N_2)(\epsilon + 1)s_1^* - 2\epsilon \right) q_1^*}{N_1 + N_2 - \epsilon}.$$

We can sign the latter derivative as follows:

$$\frac{\partial \pi_1}{\partial t} \propto (\epsilon + 1)(N_1 + N_2)(\epsilon(c_1 - t)) - N_2 \Delta) - 2\epsilon^2 \left(N_1(c_1 - t) + N_2(c_2 - t) \right)$$

$$= - \left((\epsilon + 1)(N_1 + N_2)(N_2 - \epsilon) + 2\epsilon^2 N_1 \right) \Delta + (\epsilon(1 - \epsilon)(N_1 + N_2))(c_2 - t)$$

$$\Rightarrow \frac{\partial \pi_1}{\partial t} > 0 \iff t < c_2 - \frac{\left((\epsilon + 1)(N_2 - \epsilon) + 2\epsilon^2 N_1 \right) \Delta}{\epsilon(1 - \epsilon)}.$$

Additionally, we can show that

$$
\begin{aligned}
\frac{\partial^2 \pi_1^*}{\partial t \partial c_2} &= \left(\frac{2\epsilon N_2}{(\epsilon(c_1 - t) - N_2 \Delta)^2} - \frac{\epsilon + 1}{\epsilon^2} \frac{(N_1 + N_2)N_2}{(N_1(c_1 - t) + N_2(c_2 - t))^2} \right) \pi_1^* \\
&\quad + \left(\frac{-2\epsilon}{\epsilon(c_1 - t) - N_2 \Delta} + \frac{\epsilon + 1}{\epsilon} \frac{N_1 + N_2}{N_1(c_1 - t) + N_2(c_2 - t)} \right) \times \\
&\quad \left(\frac{2N_2}{\epsilon(c_1 - t) - N_2 \Delta} - \frac{\epsilon + 1}{\epsilon} \frac{N_2}{N_1(c_1 - t) + N_2(c_2 - t)} \right) \pi_1^* \\
&= -N_2 \left(\frac{(\epsilon + 1)(N_1 + N_2)s_1^{*2} - 2\epsilon}{(\epsilon s_1^* p^*(N_1 + N_2 - \epsilon))^2} \right) \frac{\epsilon s_1^*(p^*(N_1 + N_2 - \epsilon))}{(\epsilon + 1)(N_1 + N_2)s_1^* - 2\epsilon} \frac{\partial \pi_1^*}{\partial t} \\
&\quad + \left(\frac{2N_2}{\epsilon s_1^* p^*(N_1 + N_2 - \epsilon)} - \frac{\epsilon + 1}{\epsilon} \frac{N_2}{p^*(N_1 + N_2 - \epsilon)} \right) \frac{\partial \pi_1^*}{\partial t} \\
&\propto \left(2 - (\epsilon + 1)s_1^* - \frac{(\epsilon + 1)(N_1 + N_2)s_1^{*2} - 2\epsilon}{(\epsilon + 1)(N_1 + N_2)s_1^* - 2\epsilon} \right) N_2 \frac{\partial \pi_1^*}{\partial t}.
\end{aligned}
$$

The sign of the above is ambiguous. However, note that

$$
\frac{\partial \pi_1^*}{\partial t} = ((N_1 + N_2)(\epsilon + 1)s_1^* - 2\epsilon) \frac{N_2}{N_1 + N_2 - \epsilon} q_1^* \quad \Rightarrow \quad \frac{\partial \pi_1^*}{\partial t} \propto ((N_1 + N_2)(\epsilon + 1)s_1^* - 2\epsilon).
$$

It follows that

$$
\frac{\partial^2 \pi_1^*}{\partial t \partial c_2} \propto (2 - (\epsilon + 1)s_1^*)((N_1 + N_2)(\epsilon + 1)s_1^* - 2\epsilon) - ((N_1 + N_2)(\epsilon + 1)s_1^{*2} - 2\epsilon),
$$

$$
\text{and}
$$

$$
\frac{\partial^2 \pi_1^*}{\partial t \partial c_2} < 0 \iff ((N_1 + N_2)(\epsilon + 1)s_1^* - 2\epsilon) < \underbrace{\frac{2\epsilon(s_1^* - 1)}{2 - (\epsilon + 2)s_1^*}}_{<0}
$$

where in the last step we used that $N_1 > N_2 \geq 1 \Rightarrow N_1 + N_2 > 3$, and that $s_1^* < s_2^*$. Combining these two properties of the model, we have that $s_1^* < \frac{1}{3}$, and consequently

$$
\frac{2\epsilon(s_1^* - 1)}{2 - (\epsilon + 2)s_1^*} < 0.
$$

Thus we can conclude that whenever $\frac{\partial \pi_1^*}{\partial t} > 0$ (implying that $((N_1 + N_2)(\epsilon + 1)s_1^* - 2\epsilon) > 0$), we always have that $\frac{\partial^2 \pi_1^*}{\partial t \partial c_2} > 0$.

In the above calculations, we have used that

$$q_1^* = \frac{\epsilon(c_1 - t) - N_2\Delta}{a\epsilon(N_1 + N_2 - \epsilon)}Q^{*\epsilon+1} \Rightarrow \epsilon(c_1 - t) - N_2\Delta = s_1^* aQ^{*-\epsilon}\epsilon(N_1 + N_2 - \epsilon) = \epsilon s_1^* p^*(N_1 + N_2 - \epsilon),$$

$$p^* = \frac{N_1(c_1 - t) + N_2(c_2 - t)}{N_1 + N_2 - \epsilon} \Rightarrow N_1(c_1 - t) + N_2(c_2 - t) = p^*(N_1 + N_2 - \epsilon),$$

$$\frac{\partial \pi_1^*}{\partial t} = \left(\frac{(\epsilon + 1)(N_1 + N_2)}{\epsilon\left(p^*(N_1 + N_2 - \epsilon)\right)} - \frac{2\epsilon}{\epsilon s_1^* p^*(N_1 + N_2 - \epsilon)} \right) \pi_1^* \Rightarrow \pi_1^* = \frac{\epsilon s_1^*\left(p^*(N_1 + N_2 - \epsilon)\right)}{(\epsilon + 1)(N_1 + N_2)s_1^* - 2\epsilon}\frac{\partial \pi_1^*}{\partial t}$$

We have also that

$$\frac{\partial \pi_2}{\partial t} = \frac{\left((N_1 + N_2)(\epsilon + 1)s_2^* - 2\epsilon\right)q_2^*}{N_1 + N_2 - \epsilon} > 0.$$

We can sign the expression as follows:

$$\begin{aligned}
\frac{\partial \pi_2^*}{\partial t} &\propto (\epsilon + 1)(N_1 + N_2)(\epsilon(c_2 - t)) + N_1\Delta) - 2\epsilon^2 N_1(c_1 - t) + N_2(c_2 - t)) \\
&= \left[(\epsilon + 1)(N_1 + N_2) - 2\epsilon^2\right]N_1(c_1 - t) + \left[(\epsilon + 1)(N_1 + \epsilon - 2\epsilon^2)N_2 + (\epsilon + 1)(N_1 + \epsilon)N_1\right](c_2 - t) > 0
\end{aligned}$$

where we have used that

$$(\epsilon + 1)(N_1 + N_2) > 3 \text{ and } 2\epsilon^2 < 2 \Rightarrow (\epsilon + 1)(N_1 + N_2) - 2\epsilon^2 > 0,$$
$$N_1 \geq 2 \Rightarrow N_1 + \epsilon - 2\epsilon^2 > 0.$$

Proposition 1.c. *In the linear ad valorem model, the equilibrium profit of firm i is given by $\pi_1^* = bq_i^{*2}$,*

$$\frac{\partial \pi_1^*}{\partial \alpha} = \frac{N_2\Delta + c_1}{N_1 + N_2 + 1}2q_1^* > 0,$$

while

$$\frac{\partial \pi_2^*}{\partial \alpha} = \frac{-N_1\Delta + c_2}{N_1 + N_2 + 1}2q_2^* > 0 \iff c_2 > N_1\Delta.$$

The cross-derivative of the high-cost firm's profit in equilibrium with respect to α and c_2 is positive whenever N_2 is large enough:

$$\frac{\partial^2 \pi_1^*}{\partial \alpha \partial c_2} = \frac{-2N_2\left(a - 2(1 - \alpha)(N_2 + 1)c_1 - 2(1 - \alpha)N_2 c_2\right)}{b\left(N_1 + N_2 + 1\right)^2},$$
$$\frac{\partial^2 \pi_1^*}{\partial \alpha \partial c_2} > 0 \iff \frac{a - (2(1 - \alpha)c_1)}{2(1 - \alpha)(c_1 - c_2)} < N_2.$$

In the case of an additional $1 - \beta$ factor decrease in the cost of low-cost firms, we have that the cross-derivative $\frac{\partial^2 \pi_1}{\partial \alpha \partial c_2}$ is positive if β is large enough:

$$\frac{\partial^2 \pi_1^*}{\partial \alpha \partial c_2} = \frac{-2N_2(1-\beta)\left(a - (2(1-\alpha)(N_2+1)c_1 - 2(1-\alpha)N_2(1-\beta)c_2\right)}{b\left(N_1+N_2+1\right)^2},$$

$$\frac{\partial^2 \pi_1^*}{\partial \alpha \partial c_2} > 0 \iff \frac{1}{c_2}\left(\frac{a - 2(1-\alpha)c_1}{2(1-\alpha)N_2} - (c_1 - c_2)\right) < \beta.$$

Proof. We have that equilibrium quantities are given as

$$q_1^* = \frac{a - (N_2+1)(1-\alpha)c_1 + N_2(1-\alpha)c_2}{b(N_1+N_2+1)}$$

$$q_2^* = \frac{a + N_1(1-\alpha)c_1 - (N_1+1)(1-\alpha)c_2}{b(N_1+N_2+1)},$$

and the total quantity produced is

$$Q^* = N_1 q_1^* + N_2 q_2^* = \frac{(N_1+N_2)a - N_1(1-\alpha)c_1 - N_2(1-\alpha)c_2}{b(N_1+N_2+1)}.$$

The equilirbium price is given by

$$p^* = \frac{a + N_1(1-\alpha)c_1 + N_2(1-\alpha)c_2}{N_1+N_2+1}.$$

The expression of equlirbrium quilibrium profits follow from the FOC, $\pi_i^* = bq_i^{*2}$. We obtain

$$\frac{\partial \pi_1^*}{\partial \alpha} = 2bq_1^* \frac{\partial q_1^*}{\partial \alpha} = \frac{N_2\Delta + c_1}{N_1+N_2+1} 2q_1^* > 0,$$
$$\text{and}$$
$$\frac{\partial \pi_2^*}{\partial \alpha} = 2bq_2^* \frac{\partial q_2^*}{\partial \alpha} = \frac{-N_1\Delta + c_2}{N_1+N_2+1} 2q_2^* > 0 \iff c_2 > N_1\Delta.$$

It follows that $\frac{\partial^2 \pi_1}{\partial \alpha \partial c_2} > 0$ whenever N_2 is large enough. Note that we have $\frac{\partial^2 q_1^*}{\partial \alpha \partial c_2} = \frac{-N_2}{b(N_1+N_2+1)} < 0$

If there is a reduction of marginal costs of large firms by a factor $1 - \beta$, we have the following. The equilibrium quantity for type 1 firm is given by

$$q_1^* = \frac{a - (N_2+1)(1-\alpha)c_1 + N_2(1-\alpha)(1-\beta)c_2}{b(N_1+N_2+1)}$$

and

$$\frac{\partial \pi_1^*}{\partial \alpha} = \frac{2\left((N_2+1)c_1 - N_2(1-\beta)c_2\right)}{N_1+N_2+1} q_1^* > 0$$
$$\text{and}$$
$$\frac{\partial^2 \pi_1^*}{\partial \alpha \partial c_2} = \frac{2\left((N_2+1)c_1 - N_2(1-\beta)c_2\right)}{N_1+N_2+1} \frac{\partial q_1^*}{\partial c_2} + \frac{-2N_2(1-\beta)}{b(N_1+N_2+1)} q_1^*.$$

The expressions for the cross derivative follow and the sign is ambiguous, but $\frac{\partial^2 \pi_1}{\partial \alpha \partial c_2} > 0$ whenever β is large enough.

Proposition 1.d. *In the **linear model with unit subsidy**, the equilibrium profits of firm of type i are given by* $\pi_i^* = b q_i^{*2}$. *We have that:*

$$\frac{\partial \pi_i^*}{\partial t} = \frac{2 q_i^*}{N_1 + N_2 + 1} > 0,$$

and

$$\frac{\partial^2 \pi_1^*}{\partial t \partial c_2} = \frac{2 N_2}{b(N_1 + N_2 + 1)^2} > 0.$$

Proof. We obtain that equilibrium quantities are

$$q_1^* = \frac{a - N_2 \Delta - (c_1 - t)}{b(N_1 + N_2 + 1)}$$

$$q_2^* = \frac{a + N_1 \Delta - (c_2 - t)}{b(N_1 + N_2 + 1)}.$$

with total quantity produced

$$Q^* = N_1 q_1^* + N_2 q_2^* = \frac{(N_1 + N_2)a - N_1(c_1 - t) - N_2(c_2 - t)}{b(N_1 + N_2 + 1)}.$$

The price will be

$$p^* = \frac{a + N_1(c_1 - t) + N_2(c_2 - t)}{N_1 + N_2 + 1}$$

We find that

$$\frac{\partial \pi_i^*}{\partial t} = \frac{2 q_i^*}{N_1 + N_2 + 1} > 0.$$

Note also that $\frac{\partial q_1^*}{\partial t} = \frac{1}{b(N_1+N_2+1)} > 0$ *and* $\frac{\partial q_1^*}{\partial c_2} = \frac{N_2}{b(N_1+N_2+1)} > 0$. *It is also easy to see that* $\frac{\partial^2 \pi_1}{\partial t \partial c_2} > 0$:

$$\frac{\partial^2 \pi_1}{\partial t \partial c_2} = \frac{2}{N_1 + N_2 + 1} \frac{\partial q_1^*}{\partial c_2} = \frac{2 N_2}{b(N_1 + N_2 + 1)^2} > 0.$$

When marginal costs of large firms are reduced by a factor $1 - \beta$, *we find*

$$q_1^* = \frac{a - N_2(c_1 - (1-\beta)c_2) - (c_1 - t)}{b(N_1 + N_2 + 1)},$$

and

$$\frac{\partial \pi_1^*}{\partial t} = \frac{2 q_1^*}{N_1 + N_2 + 1} > 0,$$

and

$$\frac{\partial^2 \pi_1^*}{\partial \alpha \partial c_2} = \frac{2}{N_1 + N_2 + 1} \frac{\partial q_1^*}{\partial c_2} = \frac{2 \beta N_2}{b(N_1 + N_2 + 1)^2} > 0.$$