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The Leverage-Collusion Interplay in Product and Labor Markets: Evidence from Portuguese Administrative Microdata

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Abstract

I study how firm leverage relates to collusion across both product and labor markets, using the universe of Portuguese non-financial firms over 2010–2022. Exploiting cartel decisions by the *Autoridade da Concorrência* and the minimum wage focal-point methodology for tacit wage coordination, I trace the leverage–collusion nexus in two causal directions and across distinct collusive settings. I find that pre-entry leverage positively predicts cartel formation in both product and labor markets, with the association being stronger among product market cartel firms accumulating debt than among those reducing it. The within-firm leverage dynamics diverge sharply during active cartel participation: leverage rises during product market cartel spells, contradicting the strategic deleveraging hypothesis, while it falls significantly during labor market cartel spells, consistent with that prediction. For tacit minimum wage collusion, firms engaged in wage coordination carry persistently higher leverage. The asymmetry between product and labor market cartels is a novel result, suggesting that the financial mechanism linking leverage and collusion depends on the type of market in which coordination occurs, with heterogeneous implications for capital structure theory, competition policy, and antitrust enforcement.

Keywords: collusion, leverage, capital structure, cartels, minimum wage, labor market, Portugal

JEL Classification: D22, D43, G32, J31, K21, L41

Resumo

Nesta dissertação estudo a relação entre a alavancagem das empresas e a colusão nos mercados do produto e do trabalho, utilizando o universo de empresas não financeiras portuguesas no período 2010–2022. Com base nas decisões da Autoridade da Concorrência e a metodologia do salário mínimo como ponto focal para a coordenação tácita de salários, analiso a relação alavancagem–colusão em duas direcções causais e em contextos colusivos distintos. Concluo que a alavancagem pré-entrada prevê positivamente a formação de cartel em ambos os mercados, sendo esta associação mais forte nas empresas do mercado do produto com alavancagem crescente do que nas empresas que a reduzem. A dinâmica da alavancagem dentro da mesma empresa diverge marcadamente durante os períodos de cartel activo: a alavancagem sobe durante os cartéis no mercado do produto, contrariando a hipótese de desalavancagem estratégica, enquanto cai significativamente durante os cartéis no mercado de trabalho, em linha com essa previsão. Para a colusão tácita no salário mínimo, as empresas com sinais de coordenação salarial apresentam uma alavancagem persistentemente mais elevada. A assimetria identificada entre os dois tipos de cartel é um resultado novo, sugerindo que o mecanismo financeiro que liga alavancagem e colusão depende do tipo de mercado em que a coordenação ocorre, com implicações heterogéneas para a teoria da estrutura de capital, a política de concorrência e a aplicação do direito *antitrust*.

Palavras-chave: colusão, alavancagem, estrutura de capital, cartéis, salário mínimo, mercado de trabalho, Portugal

Classificação JEL: D22, D43, G32, J31, K21, L41

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List of Acronyms

AdC	Portuguese Competition Authority
AR	Anderson–Rubin
CAE	Portuguese Classification of Economic Activities
CEO	Chief Executive Officer
DOJ	Department of Justice
FE	Fixed Effects
GDP	Gross Domestic Product
HHI	Herfindahl-Hirschman Index
INE	Portuguese National Statistics Institute
IV	Instrumental Variable
KP	Kleibergen–Paap
LBO	Leveraged Buyout
LM	Lagrange Multiplier
MLRM	Multiple Linear Regression Model
MW	Minimum Wage
NPC_FIC	Anonymous Firm Identifier
NUTS	Nomenclature of Territorial Units for Statistics
OECD	Organization for Economic Co-operation and Development
OLS	Ordinary Least Squares
PIC	Private International Cartels
PPML	Poisson Pseudo-Maximum-Likelihood
QP	*Quadros de Pessoal*
ROA	Return on Assets
ROE	Return on Equity
SABI	*Sistema de Análise de Balanços Ibéricos*
SCIE	*Sistema de Contas Integradas de Empresas*
SMEs	Small and Medium Enterprises
SSM	Single Supervisory Mechanism
SY	Stock–Yogo

1 Introduction

Collusion between firms suppresses competition through two distinct channels. In product markets, firms coordinate prices above competitive levels, divide customers or territories, or manipulate public procurement bids, capturing rents from consumers. In labor markets, firms coordinate wages below competitive levels or agree not to recruit each other's employees through no-poach agreements, suppressing worker mobility, and transferring rents from workers to firms. The empirical literature has devoted considerably more attention to the first channel, relying predominantly on databases of convicted price-fixing cartels such as the Private International Cartels dataset of Connor (2020) and the European Commission's cartel decisions database. Labor market collusion remains comparatively understudied, particularly in the European context. In this dissertation, I examine both channels, with a focus on how firm financial structure interacts with collusive behavior and, separately, how the statutory minimum wage may function as a focal point for collusion among low-wage employers.

The starting point is theoretical. Maksimovic (1988) shows that leverage shapes the incentives to sustain collusion: when debt payments are large, equity holders discount future collusion profits relative to the immediate gain from deviating, undermining the credibility of punishment threats. A maximum debt level therefore exists above which collusion becomes individually irrational, a threshold that shifts with punishment severity and the frequency of firm interaction. The two directions of this relationship have been studied separately in the empirical literature. Ferrés et al. (2021) find that firms reduce book leverage during active cartel spells, driven by debt repayment, and concentrated in the early years of collusion. Levenstein and Suslow (2016) reveal that higher industry leverage is associated with lower rates of cartel entry, as heavily indebted firms face debt-service pressures that reduce the attractiveness of long-horizon collusive payoffs.

Building on this evidence, the empirical analysis is organized around five research hypotheses. The first two (H1 and H2) ask whether pre-entry firm leverage predicts cartel formation, treating the cartel dummy as the dependent variable and leverage as the key regressor. The third and fourth hypotheses (H3 and H4) reverse the dependent variable, asking whether active cartel membership reduces firm leverage over time. Both questions are studied for product market cartels and for labor market cartels identified through decisions of the *Autoridade da Concorrência*.

The fifth hypothesis (H5) extends the analysis to tacit coordination, based on Shelkova (2015) and Knittel and Stango (2003). When the minimum wage lies below the competitive wage (non-binding) but above the monopsony wage, it can serve as a Schelling focal point around which low-wage employers coordinate tacitly without direct communication. The empirical signature of this mechanism is the minimum wage spike: an anomalous mass of wages exactly at the statutory minimum. Using matched employer-employee data, I estimate whether this spike reflects collusive wage-setting, and if so, whether firms exhibiting signs of minimum wage collusion also carry systematically different leverage.

All three analyses use Portuguese data covering the period 2010 to 2022, combining administrative records on cartel decisions from AdC, firm balance sheets from the *Sistema de Contas Integradas de Empresas*, and worker-level wage data from the *Quadros de Pessoal*. To the best of my knowledge, this is the first study to examine these questions in the Portuguese context, where the minimum wage has risen substantially over the sample period and where cartel enforcement has intensified in recent years.

The remainder of the dissertation is organized as follows. Section 2 reviews the relevant literature. Section 3 formulates the five research hypotheses. Section 4 describes the data and presents descriptive statistics. Section 5 presents the empirical methodology. Section 6 reports the results. Section 7 discusses the findings, including their implications and limitations. Section 8 presents the conclusions.

2 Literature Review

2.1 Collusion: Theory and Stability

Collusion as an economic phenomenon was formalized by Stigler (1964), who identified the three conditions that any successful cartel must satisfy: firms must agree on a joint profit-maximizing policy, monitor compliance to detect secret deviations, and credibly threaten punishment sufficient to deter defection. These conditions are harder to satisfy when markets are fragmented, buyers are heterogeneous, and monitoring is costly. Then, the stability of collusion depends critically on the discount factor: Friedman (1971) showed that in infinitely repeated games, the threat of reverting to Nash competition forever can sustain collusion as long as firms place sufficient weight on future payoffs.

The broader literature on the economic determinants of cartel formation and stability yields mixed conclusions. Rotemberg and Saloner (1986) examine collusion under demand uncertainty and prove that collusion is harder to sustain during booms, since the temptation to deviate is greatest when current demand is high, and punishment depends on future demand that is unaffected by the current state, making collusion counter-cyclical. Haltiwanger and Harrington (1991) challenge this by introducing deterministic cyclical demand, where firms in an upturn anticipate high future profits and are therefore more reluctant to trigger a price war, yielding pro-cyclical collusion. Borenstein and Shepard (1996) test these competing predictions using California gasoline price data and reveal evidence consistent with the pro-cyclical view.

García et al. (2022) contribute to this debate using the European Commission cartel data from 1997 to 2018, finding that cartels are more likely to form when managers report positive business evolution and less likely to break up under the same conditions, with sanctions by the European Commission acting as both a deterrent to new cartels and a destabilizing force on existing ones. Lenhard (2021) shows through a Hotelling model with financial constraints that cartel stability is non-monotonic in the interest rate, producing a U-shaped relationship empirically supported by European Commission data, which implies that low-rate environments may inadvertently facilitate collusion. Hellwig and Hüschelrath (2018) study the internal dynamics of cartel breakdown using a sample of 615 firms in 114 cartels convicted by the European Commission between 1999 and 2016. They demonstrate that when one firm exits a cartel, the probability that the remaining members also exit rises substantially, suggesting that cartels are fragile once defection begins. However, high interest rates, slow individual firm exits, because higher capital costs reduce the attractiveness of the competitive outside option, making firms more willing to stay in the collusive arrangement despite the temptation to deviate. Hyytinen et al. (2018), using a hidden Markov model on Finnish legal cartels, find pro-cyclical cartel formation responding to favorable GDP shocks. Levenstein and Suslow (2011) notice that firm-specific impatience rather than business cycles drives cartel duration, while Paha (2017) shows that negative demand shocks, that create excess productive capacity, increase the incremental value of collusion, particularly when capacities depreciate slowly, as firms face prolonged periods of underutilized assets and have stronger incentives to coordinate prices rather than compete them down.

2.2 Capital Structure and Collusion

Cartel agreements are difficult to enforce, since they are illegal. This makes them potentially unstable, because each cartel firm has an incentive to deviate from the agreement and steal profits from the other cartel firms. Maksimovic (1988) analyzes how financial leverage affects this incentive to deviate, exclusively focused on the product market. In his model, which takes the form of a symmetric Bertrand super game, cartel firms use grim-trigger strategies (Friedman, 1971) to enforce a collusion agreement: any deviation is punished by having all cartel firms revert to unfettered competition, and the threat of losing future collusion profits keeps cartel firms from deviating. In this setting, the critical discount factor required to sustain cooperation equals to $(n-1)/n$, where n is the number of firms. Maksimovic (1988) introduces leverage into this structure by noting that debt obligations raise the firm's required rate of return on future collusion profits, shifting its effective discount factor above the critical threshold and eroding the sustainability of the cartel agreement. Denoting the collusive profit π^c, the one-period deviation profit π^d, the Nash equilibrium profit π^{nc}, the discount rate r, and the firm's debt obligation b, deviation becomes profitable when:

$$\underbrace{(\pi^c - b) + \frac{\pi^c - b}{r}}_{\text{Equity payoff: sustain collusion}} < \underbrace{(\pi^d - b) + \frac{\pi^{nc} - b}{r}}_{\text{Equity payoff: deviate}} \quad \Leftrightarrow \quad r > \frac{\pi^c - b}{\pi^d - \pi^c} \tag{1}$$

As debt obligations increase, the present value of future collusion profits shrinks relative to the immediate gain from cheating. The maximum level of debt consistent with cartel stability is $\Delta = \pi^c - r[\pi^d - \pi^c]$: beyond this threshold, the threat of punishment loses credibility, and the cartel unravels. This generates two directional predictions: (i) firms above Δ cannot sustain collusion, so high-leverage firms should be less likely to enter cartels; and (ii) firms already in cartels have an incentive to reduce leverage below Δ, or *strategically deleverage*, to preserve collusion.

Brander and Lewis (1986) provide a complementary view in which leverage operates as a strategic commitment device through limited liability. In their two-stage sequential duopoly, equity holders ignore downside states where bondholders absorb losses, leading leveraged firms to over-produce relative to the cartel optimum. Hence, leverage is structurally pro-competitive and undermines any cartel agreement that requires output restriction.

Empirical evidence on leverage and competitive behavior in general is established by Chevalier (1995) and Phillips (1995). Using LBOs in the US supermarket industry, Chevalier (1995) documents that highly leveraged chains lose market share to unleveraged rivals, who respond by entering and expanding, while prices rise when all firms are leveraged and fall when rivals are not, consistent with predatory behavior against financially constrained incumbents. Phillips (1995) extends this to a large cross-section of industries, finding that output fell and prices rose in three out of four industries following sharp leverage increases, with the exception occurring precisely where rivals had low leverage and entry was easy. Together, these studies confirm that the competitive consequences of debt depend critically on the leverage of rivals, a consideration that carries over directly to the cartel setting studied in this dissertation.

Spagnolo (2000) shows that stock-based compensation facilitates tacit collusion by linking managers' pay to the stock price, which rational markets discount immediately upon a deviation. The stock price falls in the same period as the breach, so managers internalize the future cost of triggering a price war, making deviation less attractive. Bloomfield et al. (2023) show that cartel members are around 60% more likely to use peer-based relative performance evaluation in CEO pay, and drop it after cartel busts, suggesting that governance structures adapt to support collusive coordination.

Ferrés et al. (2021), studying the effect of active cartel membership on financial structure, using the PIC database, show that book leverage falls by around 3 percentage points during active cartel spells, with the reduction concentrated in the first two years, driven by debt repayment rather than asset growth, and more pronounced in cartels with higher prior leverage and stronger competitive pressure. Levenstein and Suslow (2016), by contrast, study what predicts cartel formation and document that higher industry leverage is associated with lower rates of entry into collusion, consistent with the participation constraint: heavily indebted firms face short-term cash flow pressures that make the long-horizon payoffs from collusion insufficiently attractive relative to the immediate need to service existing debt.

Forsbacka et al. (2025) report that higher economy-wide interest rates increase cartel formation and reduce the likelihood of cartel death among Swedish legal cartels. Although higher rates compress the discount factor — lowering the value of future collusive profits — this effect is outweighed by a cost-of-capital channel: as external financing becomes more expensive, deviating from a cartel requires investment capacity that firms can no longer afford, and defection may threaten survival in a high-rate environment. Collusion thus emerges as a margin-protection strategy rather than a growth one. Huric-Larsen (2025) contributes a complementary firm-level perspective, finding that 70.6 percent of firms sanctioned by the Danish competition authorities exhibited a declining ROE trend in the period preceding cartel formation, which reversed during collusion and reverted after breakdown. This suggests that financial stress acts as a facilitating factor in cartel entry: firms may turn to collusion not merely to maximize profits but to arrest a deterioration in returns that threatens their access to capital and investor confidence.

2.3 Labor Market Collusion and the Minimum Wage as a Focal Point

The connection between collusion and capital structure extends to labor markets. Gonzaga et al. (2013) show that collusion in the labor market, under the assumption of labor as the sole input, produces outcomes equivalent to product market collusion: higher product prices, lower output, lower employment, and lower wages. This equivalence makes the Maksimovic (1988) framework, in principle, applicable to labor cartels, although the underlying mechanism may differ: product cartels generate abnormal profits that can be used to service or retire debt, whereas labor cartels generate cost savings whose implications for leverage depend on how those savings are used.

Horizontal wage collusion takes two main forms: explicit wage-fixing agreements, including no-poach pacts, where firms agree not to recruit from each other, and tacit coordination, where firms adjust wages by observing rivals without direct communication (Masur & Posner, 2023; Pecman et al., 2021).

Herrera-Caicedo et al. (2025) document that common ownership, through shared executives or board members, raises the probability of no-poach agreements by 12 percentage points relative to a baseline of 1.2%. Delabastita and Rubens (2025) provide evidence that employer collusion in labor markets suppresses wages below the competitive benchmark. Liang (2024) extends this literature to capital structure, finding that firms more exposed to the minimum wage reduce leverage by 0.5 to 0.9 percentage points, through increased cash holdings and reduced long-term debt, with effects concentrated among small firms. Arabzadeh et al. (2024) demonstrate that the minimum wage reduces within-firm wage dispersion, with the effect amplified in financially constrained firms, specifically those with leverage above 50%, reflecting a stronger wage compression at the bottom of the distribution when firms have less financial flexibility.

The second strand of literature relevant to this dissertation concerns the minimum wage as a focal point for tacit collusion. Schelling (1960) argued that when parties share common interests but cannot communicate, they coordinate on salient reference points that make one option stand out. Knittel and Stango (2003) applied this logic to credit card markets, showing that non-binding state interest rate ceilings in the United States during the 1980s acted as focal points for tacit collusion among card issuers, with the collusion probability rising with market concentration and issuer costs. Using a double-hurdle model adapted from Cragg (1971), they estimate the probability that the observed prices at the ceiling reflect coordination rather than binding regulation.

Shelkova (2015) transplants this framework to the labor market. Building on the Folk Theorem, she models an infinitely repeated wage-setting game in which any wage between the monopsony and competitive levels can be sustained with trigger strategies when firms are sufficiently patient. A non-binding minimum wage becomes a natural focal point for coordination because it is salient, legally visible, and provides the highest payoff that all low-wage employers can plausibly sustain. The empirical signature is the minimum wage spike, the excess mass of wages observed exactly at the statutory minimum. Using Current Population Survey data on US service workers from 1990 to 2002, she estimates that around 19 percent of minimum wage earners are affected by collusive wage-setting, with collusion more prevalent in metropolitan areas, among part-time employers, and in the period following federal minimum wage hikes when the new level becomes the relevant focal point. Shy and Zhang (2018) provides a complementary theoretical treatment, showing that in oligopsonistic labor markets collusion at the minimum wage is sustainable when the minimum wage lies below the competitive wage, worker mobility costs are high, and the number of workers is large. Sharma (2025) documents this mechanism empirically in the Indian garment industry, finding that collusion keeps wages and employment below the imperfectly competitive benchmark, and that minimum wage increases can raise both wages and employment further than simply eliminating collusion, because more productive firms re-coordinate at the new higher level.

2.4 Antitrust Enforcement

Competition authorities have substantially intensified enforcement over the past three decades. The OECD Council's 1998 recommendation on hard-core cartels prompted widespread criminalization, expanded private litigation, and the consolidation of leniency programs. Chen and Rey (2013) characterize the optimal design of such programs, showing that some leniency is always desirable before an investigation is launched, that restricting eligibility to the first informant is optimal, and that withdrawing leniency from repeat offenders is counterproductive, as it renders the program ineffective. Between 1990 and 2025, the European Commission imposed €30.6 billion in cartel fines (adjusted for Court judgments) across 162 decisions involving 953 undertakings. Enforcement intensity peaked between 2005 and 2019, with fines of €7.8 billion in 2005–2009, €7.6 billion in 2010–2014, and €8.2 billion in 2015–2019, driven in part by landmark cases such as the Trucks cartel and the Forex spot trading cartel. The most recent period has been marked by sharp cyclical variation: fines fell to €88.9 million in 2023 and €48.7 million in 2024, the lowest levels in decades, before rebounding to €859 million in 2025 [1].

Miller (2009) offers the first independent empirical evaluation of the US leniency program introduced in 1993, applying a Markov model of cartel behavior to the complete set of DOJ indictments between 1985 and 2005. The results show that cartel discoveries increased immediately after the program was introduced and subsequently fell below pre-leniency lev-

[1] Cartel statistics: https://competition-policy.ec.europa.eu/antitrust-and-cartels/cartels-cases-and-statistics_en.

els. The European Commission revised its own leniency program along similar lines in 2002, extending automatic amnesty to the first confessor. Despite this convergence, OECD (2023) documents that leniency applications across OECD jurisdictions fell by 58% between 2015 and 2021, raising questions about the long-run sustainability of leniency as the primary detection tool.

The reliability of cartel data poses an additional challenge. Harrington and Wei (2017) caution that cartel data based on discovered cases may be subject to selection bias, since cartels with heterogeneous collapse and discovery probabilities produce discovered samples that over-represent stable and easily detected cases. Gerlach and Li (2024) document that hard-core cartels persist despite stronger enforcement, consistent with the view that collusion remains profitable even at elevated detection probabilities.

In terms of deterrence, the literature has struggled with a fundamental identification problem: estimating the probability of detection requires information on cartels that were never discovered. Combe and Monnier (2011) analyze 64 European cartel decisions between 1975 and 2009 and find that, assuming a 15% probability of detection, only one fine falls within the dissuasive range and none exceeds it, concluding that Commission fines are systematically below the level required to deter cartel formation. Katsoulacos and Ulph (2013) extend this analysis by introducing considerations relating to the timing of penalty decisions, showing that the optimal penalty is approximately 75% of that implied by the conventional formula. This leads them to conclude, in contrast to Combe and Monnier (2011), that existing European fines are broadly within the range supported by calculations of optimal penalties.

3 Hypothesis Formulation

Drawing on the literature previously reviewed, the empirical analysis is organized around five research hypotheses. The first two concern the direction from leverage to cartel formation. The Maksimovic (1988) participation constraint predicts that more leveraged firms should be less likely to form cartels, as the debt burden reduces the relative attractiveness of long-horizon collusion payoffs. However, alternative mechanisms predict the opposite sign: financially stressed firms may seek collusion precisely to stabilize revenues and service existing debt Huric-Larsen (2025), and firms accumulating new liabilities may have the strongest incentive to lock in revenue streams through coordination (Brander & Lewis, 1986).

H1 (Product cartel): *Pre-entry firm leverage predicts product market cartel formation.*

H2 (Labor cartel): *Pre-entry firm leverage predicts labor market cartel formation.*

The sign of both hypotheses is treated as an empirical question, with the Maksimovic (1988) participation constraint predicting a negative coefficient and the financial stress mechanism predicting a positive one.

The third and fourth hypotheses concern the reverse causality: whether active cartel membership affects leverage within the same firm. These hypotheses are consistent with the strategic deleveraging mechanism proposed by Maksimovic (1988) and empirically supported by Ferrés et al. (2021). Despite the possibility that cartel profits may be reinvested rather than used to retire debt, the dominant prediction of the strategic deleveraging hypothesis is that collusion reduces leverage.

H3 (Product cartel effect on leverage): *Active product cartel membership is associated with a reduction in firm leverage.*

H4 (Labor cartel effect on leverage): *Active labor cartel membership is associated with a reduction in firm leverage.*

H3 and H4 are treated separately because the financial consequences of product and labor market collusion are not equivalent: product cartels raise revenues, while labor cartels reduce costs. If the additional cash flow from product cartel profits is reinvested in the firm or used to expand operations rather than to retire debt, leverage may rise even as the cartel stabilizes, contrary to the prediction of the strategic deleveraging hypothesis. Labor cartels, by contrast, reduce the wage bill directly, freeing cash flow that could be used to service or reduce debt. The contrast between H3 and H4 thus provides a natural test of whether the strategic deleveraging mechanism operates uniformly across cartel types.

The fifth hypothesis extends this logic to tacit coordination around the statutory minimum wage. If firms coordinate wages at the minimum wage, labor costs may be compressed relative to the competitive benchmark, increasing internal cash flow and allowing firms to reduce leverage (Liang, 2024; Shelkova, 2015).

H5 (Minimum wage collusion effect on leverage): *Firms exhibiting evidence of tacit minimum wage collusion are associated with a reduction in firm leverage.*

4 Data

4.1 Data Description

This study draws on three data sources covering the period 2010–2022. The first is a dataset of cartel cases in Portugal, built by web scraping all decisions published on the website of the *Autoridade da Concorrência* (AdC), the Portuguese competition authority, established in 2003 under Law 18/2003. The dataset covers collusion cases, and records for each case the start and end dates of the cartel activity as well as the start and end dates of the AdC investigation. The remaining cartel observations were matched to two administrative datasets: *Quadros de Pessoal* and *Sistema de Contas Integradas de Empresas*. The *Quadros de Pessoal* (QP) is an employer-employee matched panel collected annually, covering virtually all firms with at least one employee in Portugal, and contains detailed worker-level information including wages, hours worked, and education. The *Sistema de Contas Integradas de Empresas* (SCIE) is an annual firm-level panel collected by the *Instituto Nacional de Estatística* (INE), the Portuguese national statistics office, containing balance sheet and income statement information for the universe of Portuguese non-financial firms. To ensure consistency with the SCIE: foreign firms, sectoral associations, professional orders, trade unions, and firms operating in banking, insurance, and holding companies were dropped from the AdC dataset.

Since both the QP and the SCIE anonymize firms through an internal identifier (NPC_FIC) rather than firm names, a dedicated identification procedure was implemented to link the AdC cartel firms, identified by name, to their anonymous counterparts in the administrative data. The matching was performed using SABI, a commercial database containing named firm-level financial information for Portuguese companies. Since some cartel firms were active before or after the sample period, or changed status over time, the matching was conducted for multiple reference years (2011, 2017, and 2022) to maximize coverage across the full panel.

For the SCIE, the matching exploits the joint uniqueness of three balance sheet variables, namely total assets, equity, and total liabilities, combined with the NUTS-2 region when available. Cases with multiple candidates were flagged as ambiguous and excluded.

For the QP, the matching follows a sequential approach using turnover, share capital, and two-digit CAE sector code. Cases yielding multiple candidates were likewise excluded.

4.2 Descriptive Statistics

The firms' panel, described in Table 1, covers 15,829,530 firm-year observations corresponding to 3,007,418 unique firms over the period 2010–2022.

Table 1: Panel Overview

Statistic	Value
Total observations	15,829,530
Unique firms	3,007,418
Years	2010–2022
Number of years	13

Drawing on the case descriptions published by the AdC, the enforcement record is classified into three categories: product market collusion, labor market collusion, and abuse of dominant position. Product market collusion refers to agreements or concerted practices between firms that restrict competition in the markets where they sell goods or services, in its most direct form involving competitors coordinating on prices, dividing customers or territories, or manipulating bids in public procurement. Labor market collusion refers to agreements between competing employers that restrict rivalry on the demand side of the labor market, where firms coordinate on wages offered to workers or, more commonly, on recruitment practices, suppressing worker mobility and keeping wages below the level that would prevail under competition. Abuse of dominant position is a unilateral practice in which a single firm exploits its market power to impose unfair conditions, foreclose rivals, or discriminate against trading partners. Given its unilateral nature, this category falls outside the scope of the analysis and is excluded from the main specifications. Product market collusion is the dominant infringement type, with 348 firm-year observations flagged as cartel-active and 80 unique firms affected at some point during the sample, as reported in Table 2. Labor market collusion affects 35 firms across 114 firm-year observations.

Table 2: Cartel Dummies - Global Frequencies

Variable	Obs. = 1	Unique Firms
Product Market Collusion	348	80
Labor Market Collusion	114	35

The AdC sample contains 34 active cartels in product market at some point between 2010 and 2022, as reported in Table 3. The sectoral distribution of the 125 firms investigated by the AdC is reported in Table 4, where Wholesale & Retail Trade are 17.6% (including major supermarket chains and beverage suppliers involved in hub-and-spoke price-fixing arrangements), Professional Services are 16.8% (covering telerradiology providers and business information database firms), Manufacturing are 14.4% (comprising producers of foam, envelopes, and cable), and Construction are 4.8% (covering railway maintenance and pre-fabricated school structures). The legal basis for all product market cases is Article 9 of Law 19/2012 and, where trade between member states is affected, Article 101 of the Treaty on the Functioning of the European Union.

Table 3: Summary of Cartel Cases in Portugal

	Collusion	
	Product	**Labor**
Cartels active in period	34	2
Cartels formed	17	2
Cartels ended	33	2

The most prevalent form of labor market collusion documented by the AdC is the no-poach agreement, under which firms commit not to recruit or proactively approach one another's employees, thereby eliminating the competitive pressure that would otherwise require employers to offer improved conditions to attract and retain workers. The dataset contains 2 labor market collusion cases (Table 3), with Sports & Recreation accounts for 22.4% of the 125 firms investigated by the AdC in Table 4 (reflecting the scale of the football no-poach case), and Information & Technology are 4.8% (covering the technology consulting firms involved in the no-poach agreement). These two cases represent the first labor market collusion decisions issued in Portugal, and the AdC has made explicit in both that this conduct is a cartel infringement of the same seriousness as product market collusion, subject to the same legal framework under Article 9 of Law 19/2012.

The descriptive statistics reported in Table 5 show considerable dispersion across firms. Leverage, measured by the debt-to-assets ratio, has a mean of 8.68, though the median firm sits at 0.73, indicating a heavily right-skewed distribution driven by a small number of highly leveraged observations. The extreme values in the upper tail largely reflect firms reporting very low asset bases, which mechanically generate unusually large debt-to-assets ratios. To

Table 4: Sectoral Distribution of Cartels in Portugal

Industry	**Unique firms**	**Percentage**
Manufacturing	18	14.40
Energy & Utilities	8	6.40
Construction	6	4.80
Wholesale & Retail Trade	22	17.60
Transportation & Communications	4	3.20
Information & Technology	6	4.80
Professional Services	21	16.80
Rental & Leasing Activities	4	3.20
Health	8	6.40
Sports & Recreation	28	22.40
Total	**125**	**100.00**

Notes: Manufacturing (CAE 10, 11, 17, 18, 19, 22, 24, 25, 27, 29, 30); Energy & Utilities (CAE 35, 38); Construction (CAE 41, 42); Wholesale & Retail Trade (CAE 45, 46, 47); Transportation & Communications (CAE 53, 60, 61); Information & Technology (CAE 58, 62); Professional Services (CAE 70, 73, 82); Rental & Leasing Activities (CAE 77); Health (CAE 86); Arts, Sports & Recreation (CAE 93).

mitigate the influence of these observations, leverage is winsorized at the 1st and 99th percentiles in all regression specifications. Return on assets displays a negative mean of −1.69, with the median close to zero (0.006), consistent with a sample that includes a large share of small firms reporting losses in any given year. The Herfindahl-Hirschman Index has a mean of 0.008, suggesting that most firms operate in relatively unconcentrated sectors, though the distribution has a long right tail reaching values close to 1. The median firm has a single employee and sales of 9,389 euros, reflecting the predominance of micro-firms in the Portuguese economy.

Table 5: Descriptive Statistics

Variable	**Obs.**	**Mean**	**Std. Dev.**	**Median**	**Min.**	**Max.**
Leverage	5,154,090	8.678	1,196.024	0.725	−26,839	1,956,007
ROA	5,154,090	−1.686	236.644	0.006	−261,572	35,223
Sales	15,829,530	305,789	13,100,000	9,389	0	13,400,000,000
Employees	15,829,530	3.154	49.281	1	1	26,859
HHI	15,829,530	0.008	0.020	0.004	0.001	0.936

Table 6 documents the incidence of minimum wage employment across the sample. Over the full sample period, 5,443,341 worker-year observations correspond to workers earning exactly the Portugal's statutory minimum wage (i.e., *Retribuição Mínima Mensal Garantida*), representing 13.99% of all observations. The wage gap between the two groups is substantial:

minimum wage workers earn on average €570.05 in base wages, compared to €878.12 for those above the floor, implying that above-minimum workers earn roughly 54% more. The minimum wage is highly visible, legally mandatory, and publicly announced in advance, making it a natural Schelling focal point for low-wage employers who may wish to coordinate wages tacitly.

Table 6: Minimum Wage Workers - Overview

Statistic	**Value**
Total workers at minimum wage	5,443,341
% of total observations	13.99%
Mean base wage – at minimum wage	570.05
Mean base wage – above minimum wage	878.12

Table 7 shows the five industries with the highest concentration of minimum wage workers over 2010–2022. Retail trade ranks first, with 647,685 worker-year observations at the minimum wage (11.9% of the total), followed by food and beverage services with 582,930 (10.7%), and manufacture of wearing apparel with 377,869 (6.9%). Wholesale trade and construction complete the top five. Together, these five industries account for 38.6% of all minimum wage observations, indicating that minimum wage employment is concentrated in a small number of labor-intensive sectors.

Table 7: Top 5 Industries by Minimum Wage Workers

CAE	**Industry**	**Workers at MW**	**% of total MW**
47	Retail trade, except motor vehicles	647,685	11.90
56	Food and beverage service activities	582,930	10.71
14	Manufacture of wearing apparel	377,869	6.95
46	Wholesale trade, except motor vehicles	262,664	4.83
41	Construction of buildings	231,188	4.25
Top 5 Total		**2,102,336**	**38.63**
Grand Total		**5,443,341**	**100.00**

5 Methodology

The analysis is organized around three empirical questions. First, whether pre-entry leverage predicts cartel formation. Second, whether active cartel membership affects firm leverage during the collusive spell. Third, whether the same leverage dynamics are present in the context of tacit labor market collusion, identified through the latent wage methodology of Shelkova (2015). Throughout the first two parts, cartel participation is identified through decisions published by the *Autoridade da Concorrência*, covering both product and labor market cartels.

5.1 Leverage Predicting Cartel Formation (H1, H2)

The first set of results examines whether a firm's leverage predicts its likelihood of engaging in cartel formation, for both product and labor market cartels. The cartel indicator serves as the dependent variable, with lagged leverage, measured before cartel entry, as the main regressor, testing whether pre-collusion indebtedness influences the probability of joining a cartel (H1 and H2). The baseline specification estimates this relationship at the firm-year level:

$$\text{Cartel}_{it} = \beta \,\text{Leverage}_{i,t-1} + \gamma X_{i,t-1} + \alpha_s + \delta_t + \varepsilon_{it}, \tag{2}$$

where Cartel_{it} is a dummy variable equal to one if the firm is involved in active cartel participation — either product or labor market — $\text{Leverage}_{i,t-1}$ is lagged leverage measured as debt to assets, $X_{i,t-1}$ is a matrix of lagged firm-level controls (i.e., ROA, log sales, log employees, and HHI), α_s are sector fixed effects at the two-digit CAE level, and δ_t are year fixed effects. Year fixed effects are included to absorb aggregate time-varying shocks, such as macroeconomic conditions or changes in enforcement intensity, that could simultaneously affect firm leverage and the propensity for collusion. Sector fixed effects, defined at the two-digit CAE level, control for time-invariant industry characteristics, such as barriers to entry, product homogeneity, or regulatory environment, that make collusion structurally more or less viable across sectors. Firm fixed effects are excluded from the cartel formation analysis, as the dependent variable is a rare binary event occurring in fewer than 0.01% of firm-year observations, leaving insufficient within-firm variation for identification.

An asymmetric extension decomposes leverage into increasing and decreasing spells to test whether it is the accumulation or the reduction of debt that drives the association with cartel entry:

$$\text{Cartel}_{it} = \beta^+ \left(\text{Leverage}_{i,t-1} \times \mathbf{1}^+_{i,t-1}\right) + \beta^- \left(\text{Leverage}_{i,t-1} \times \mathbf{1}^-_{i,t-1}\right) + \gamma X_{i,t-1} + \alpha_s + \delta_t + \varepsilon_{it}, \tag{3}$$

where $\mathbf{1}^+_{i,t-1}$ is a dummy equal to one if the firm recorded a positive year-on-year change in leverage in the previous period, and $\mathbf{1}^-_{i,t-1}$ is defined analogously for negative changes. β^+ and β^- capture the effect of leverage separately for firms accumulating and reducing debt, and the hypothesis of interest is whether $\beta^+ = \beta^-$.

To address residual endogeneity, the OLS estimates may be subject to bias if firms simultaneously adjust their leverage and their propensity to collude in response to common shocks, or if firms strategically adjust their capital structure in anticipation of cartel formation. Bartik shift-share instruments are therefore constructed, following Bartik (1991) and Goldsmith-Pinkham et al. (2020). The second-stage equation takes the form:

$$\text{Cartel}_{it} = \beta\, \widehat{\text{Leverage}}_{i,t-1} + \gamma X_{i,t-1} + \alpha_s + \delta_t + \varepsilon_{it}, \tag{4}$$

where $\widehat{\text{Leverage}}_{i,t-1}$ is the predicted value from the first stage. The instrument is constructed by interacting each firm's lagged employment share within its two-digit CAE sector with the economy-wide average leverage excluding the firm's own sector, so that the aggregate shift is not contaminated by sector-specific shocks:

$$Z^{\text{emp}}_{i,t-1} = s^{\text{emp}}_{i,t-1} \left(\frac{\overline{\text{Lev}}_{t-1} - \omega_{s,t-1}\, \overline{\text{Lev}}_{s,t-1}}{1 - \omega_{s,t-1}} \right) \tag{5}$$

where $\overline{\text{Lev}}_{t-1}$ is the economy-wide mean leverage, $\overline{\text{Lev}}_{s,t-1}$ is the mean leverage of sector s, $\omega_{s,t-1}$ is the sector's weight in total firms, and $s^{\text{emp}}_{i,t-1}$ is the firm's employment share within its sector. The term in parentheses corresponds to a leave-one-sector-out measure of aggregate leverage.

The second instrument exploits within-industry variation in leverage combined with firm-level exposure based on sales shares. It is formally defined as:

$$Z^{\text{sales}}_{i,t-1} = s^{\text{sales}}_{i,t-1}\, \overline{\text{Lev}}^{-i}_{s,t-1} \tag{6}$$

where $s^{\text{sales}}_{i,t-1}$ denotes firm i's share of total sales within sector s, and $\overline{\text{Lev}}^{-i}_{s,t-1}$ is the leave-one-out average leverage in sector s, excluding firm i. These instruments predict firm-level leverage using exogenous aggregate and sectoral financial conditions weighted by predetermined firm-specific exposure shares.

Instrument relevance is initially assessed using the Kleibergen–Paap rk LM statistic for underidentification in the baseline specification contemplating the endogenous regressor substituted by a single instrument. I also test the exogeneity of leverage using a standard endogeneity test to verify whether instrumentation is required. Conditional on evidence of endogeneity, I evaluate instrument strength using the Kleibergen–Paap rk F-statistic, benchmarked against the Stock–Yogo 10% maximal IV size critical value. Robustness to weak identification is further assessed with the Anderson–Rubin and Stock–Wright tests, which test the null that leverage has no effect on cartel formation while remaining valid regardless of instrument strength.

In a further specification, I augment the model with a second shift-share instrument—the one based on the firm's lagged sales share—serving as a robustness check. In this overidentified specification, instrument validity is assessed using the Hansen J test. The identifying assumption underlying the approach is that economy-wide and sectoral leverage shocks affect cartel formation only through their impact on firm-level leverage, ensuring relevance while maintaining exogeneity with respect to the error term in the cartel formation equation.

5.2 Active Cartel Membership and Leverage (H3, H4)

The second direction reverses the roles of the variables: leverage becomes the dependent variable, and active cartel membership is the main regressor, addressing H3 and H4. The question is what happens to firm leverage once a cartel is already underway. Maksimovic (1988) predicts that firms should strategically reduce their leverage during active collusion to sustain the agreement, since high debt payments make deviation more attractive to equity holders by eroding the value of the future collusive stream relative to the immediate gain from cheating. The specification takes the form:

$$\text{Leverage}_{it} = \beta\,\text{Cartel}_{it} + \gamma X_{i,t-1} + \mu_i + \alpha_s + \delta_t + (\alpha_s \times \delta_t) + \varepsilon_{it}, \tag{7}$$

where μ_i are firm fixed effects that absorb all time-invariant firm-level heterogeneity, and

$\alpha_s \times \delta_t$ are sector×year fixed effects that control for shocks specific to each sector in each year. Four specifications are estimated, progressively enriching the fixed-effect structure from sector and year alone to the most demanding combination of firm and sector×year fixed effects. The rich fixed-effects structure, and specifically the inclusion of firm fixed effects, is critical for identification: without it, the coefficient on the cartel indicator would reflect self-selection bias—namely, that firms joining cartels systematically carry lower leverage than their peers in the same sector and year—rather than the within-firm effect of cartel membership.

Additionally, potential pre-entry dynamics and post-entry leverage adjustments are assessed through a dynamic time fixed effects specification that includes two leads and two lags of the cartel indicator, estimated with firm and sector×year fixed effects.

$$\text{Leverage}_{it} = \sum_{\tau=-2}^{+2} \delta_\tau \, D_{i,t+\tau} + \gamma X_{i,t-1} + \mu_i + (\alpha_s \times \delta_t) + \varepsilon_{it}, \tag{8}$$

where $D_{i,t+\tau}$ is an indicator equal to one if firm i enters a cartel τ periods in the future (i.e., leads) or entered τ periods in the past (i.e., lags), and the period immediately before cartel onset ($\tau = -1$) is the omitted category. A linear combination test of pre-entry and post-entry coefficients helps identify whether cartel formation is responsible for statistically significant changes in leverage.

5.3 Minimum Wage Collusion and Leverage (H5)

The final part follows Shelkova (2015) to identify minimum wage collusion. The empirical signature is an anomalous concentration of wages exactly at the statutory minimum among workers whose observable characteristics would justify a higher wage.

Identification proceeds in three steps. First, a latent wage equation is estimated on workers with wages strictly above the minimum to avoid contaminating the coefficients with censoring from below (Shelkova, 2015):

$$\ln w_{jt} = \alpha + \beta_1 \, \text{age}_{jt} + \beta_2 \, \text{age}_{jt}^2 + \sum_{k=1}^{8} \delta_k \, \mathbf{1}[\text{habil}_{jt} = k] + \beta_3 \, \text{male}_{jt} + \beta_4 \, \text{parttime}_{jt} + \theta_s + \delta_t + \varepsilon_{jt}, \tag{9}$$

where w_{jt} is the base monthly wage of worker j in year t, age_{jt} is the worker's age and age_{jt}^2 its square, included as a proxy for labor market experience, habil_{jt} is their education level across eight categories, male_{jt} is an indicator for male workers, parttime_{jt} is an indicator for part-time employment, and two-digit CAE sector and year fixed effects are absorbed.

Second, the estimated coefficients are applied to all workers to produce the latent wage $\hat{w}_{jt}$, the wage each worker would receive based on their observable characteristics in the absence of any distortion at the minimum. Third, a worker is classified as potentially affected by collusive wage-setting if their observed wage equals the statutory minimum but their latent wage exceeds it. The firm-level indicator $\text{Collusion}_{it}^{MW}$ equals one if the firm has at least one such worker in year t. With this indicator in hand, the effect of minimum wage collusion on leverage was estimated:

$$\text{Leverage}_{it} = \beta\,\text{Collusion}_{it}^{MW} + \gamma X_{i,t-1} + \mu_i + \alpha_s + \delta_t + (\alpha_s \times \delta_t) + \varepsilon_{it}, \qquad (10)$$

where $\text{Collusion}_{it}^{MW}$ is the binary indicator constructed above, $X_{i,t-1}$ is the same matrix of lagged firm-level controls as before, μ_i are firm fixed effects, and $\alpha_s \times \delta_t$ are sector×year fixed effects. The four specifications differ in which subset of $\{\alpha_s, \delta_t, \mu_i, \alpha_s \times \delta_t\}$ is included, following the same structure as equation (7).

The section closes with a robustness analysis comparing three alternative operationalizations of the collusion indicator to assess the sensitivity of the results to the definition of the threshold. The first alternative replaces the binary indicator with the continuous share of affected workers in total firm employment, defined as the ratio of workers who earn exactly the statutory minimum, but have a predicted latent wage above it, to the total number of workers in the firm. This continuous measure captures the intensity of wage suppression within the firm rather than treating all firms with at least one affected worker as equivalent, and allows the estimated effect on leverage to vary with the fraction of the workforce subject to collusive wage-setting. The second alternative uses a binary indicator that requires at least ten percent of the firm's workers to satisfy the collusion criterion, imposing a more demanding threshold that guards against false positives driven by isolated prediction error in the latent wage equation. The third uses a binary indicator based on whether the share of affected workers exceeds the sample median among firms with at least one affected worker.

A dynamic time fixed effects specification similar to equation (8) is also considered to assess whether leverage exhibits significant pre-entry or post-entry adjustments.

5.4 Estimation Strategy

All multiple linear regression models (MLRMs) are estimated using OLS and PPML in parallel. PPML is motivated by the highly skewed distribution of leverage and by the rarity of collusion events in the data. While OLS remains a valid estimator, it may produce fitted values outside the admissible range for binary outcomes and can be inefficient under substantial heteroskedasticity. PPML is also supported by technical considerations: Santos Silva and Tenreyro (2006) show that it remains consistent under fairly general forms of variance misspecification. For binary dependent variables, PPML with firm fixed effects avoids the incidental parameters problem that affects nonlinear estimators such as probit and logit in large panels (e.g., Fernández-Val and Weidner (2016)), since Poisson fixed effects can be absorbed without bias Guimarães and Portugal (2010). Finally, the estimated coefficients admit a straightforward semi-elasticity interpretation. Standard errors are clustered at the firm level throughout, allowing for arbitrary heteroskedasticity and within-firm serial correlation Bertrand et al. (2004).

The control matrix ($\mathbf{X}_{i,t-1}$) includes, in all models, lagged ROA, logarithm of sales, logarithm of employees, and the Herfindahl–Hirschman index computed from sales-based market shares within each two-digit CAE sector and year. Return on assets controls for profitability, which affects both debt capacity and cartel incentives (Dasgupta & Žaldokas, 2019; Ferrés et al., 2021; Liang, 2024). The logarithm of sales and the logarithm of employees capture firm size, which determines market power and monitoring capacity (Dasgupta & Žaldokas, 2019; Ferrés et al., 2021; Liang, 2024). The Herfindahl–Hirschman index controls for market concentration, a fundamental determinant of cartel stability, since more concentrated markets facilitate the monitoring and punishment of deviations (Dasgupta & Žaldokas, 2019; Stigler, 1964). Leverage is winsorized at the 1st and 99th percentiles throughout.

6 Results

6.1 Product Market Cartel

Leverage and product market cartel formation (H1). Table 8 reports the baseline estimates for the relationship between leverage and product market collusion, as specified in equation (2). In the bivariate specification, the OLS coefficient on leverage is negative and statistically significant. Once firm-level and market controls are added, the coefficient reverses sign and becomes positive and significant, implying that the same one-million-euro increase in leverage relative to the mean is now associated with an increase of approximately 18 percentage points in the probability of collusion. The sign reversal suggests that leverage operates as a conditional instrument, with its relationship to collusion depending critically on firm size, profitability, and market structure: without controlling for these factors, leverage proxies for firm characteristics that are independently associated with lower collusion risk.

Poisson estimates broadly confirm this pattern, with the bivariate coefficient negative and significant ($\hat{\beta} \approx -0.2634$) and turning positive but insignificant once controls are introduced.

Table 8: Product Market Cartel Formation and Leverage

	OLS		Poisson (PPML)	
	(1)	(2)	(3)	(4)
Leverage$_{t-1}$	-0.0000046^{***}	0.000018^{***}	-0.2634^{***}	0.0275
	(0.000001)	(0.0000029)	(0.0757)	(0.1275)
ROA$_{t-1}$		0.000000^{***}		0.0006
		(0.000000)		(0.0010)
ln(Sales)$_{t-1}$		0.0000736^{***}		1.1664^{***}
		(0.000014)		(0.0796)
ln(Employees)$_{t-1}$		0.000196^{***}		0.0077
		(0.000044)		(0.0804)
HHI$_{t-1}$		0.0044956^{**}		6.1886^{***}
		(0.0022331)		(1.1340)
Year FE	✓	✓	✓	✓
Sector FE	✓	✓	✓	✓
Controls		✓		✓
Observations	4,371,271	3,875,211	2,067,430	1,866,653

Notes: Dependent variable is Cartel$_{it}^{PM}$, a binary indicator equal to one when the firm is engaged in active product market collusion. All specifications include year and sector (CAE-2 digit) fixed effects. Leverage$_{i,t-1}$ is measured by lagged debt-to-assets, winsorized at the 1st and 99th percentiles. Controls include ROA, log sales, log employees, and the Herfindahl–Hirschman index. Standard errors clustered at the firm level in parentheses. Poisson estimates drop singleton observations separated by fixed effects. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

The OLS estimates may be subject to endogeneity bias if firms simultaneously adjust their leverage and their propensity to collude in response to common shocks, or if firms anticipating cartel formation strategically adjust their capital structure in advance. A formal endogeneity test (table 21) rejects the null that leverage is exogenous after conditioning on firm characteristics and partialing out the controls, confirming that IV estimation is required. Table 9 addresses this endogeneity using the Bartik-style instrument described in equations (4) and (5).

The Kleibergen–Paap LM statistic rejects the null of underidentification at the 5% significance level both with and without controls (p-value $= 0.000$), indicating that the excluded instrument initially adopted is relevant and that the model is properly identified when using this shift-share instrument. However, to justify the use of one excluded instrument for one potentially endogenous variable, the adopted instrument's validity also depends on its strength. Accordingly, the Kleibergen–Paap rk Wald F-statistic exceeds the Stock–Yogo 10% maximal IV size critical value, leading to rejection of the null of weak identification and indicating that the first shift-share instrument is sufficiently strong. The Anderson–Rubin test, which is robust to weak instruments, rejects the null that leverage has no effect on product cartel formation.

In the overidentified specification, where a second instrument is added to the first-stage regression (6), the Hansen J test does not reject the null hypothesis that the instruments are jointly valid, providing no evidence against the overidentifying restrictions. However, as the endogeneity test fails to reject exogeneity in the overidentified case (Table 21, column (2)), the overidentified result is reported solely for completeness.

Across all specifications, the second-stage coefficient on leverage is positive and significant when controls are included ($\hat{\beta} \approx 0.0263$, $p < 0.01$), confirming that the positive relationship between leverage and product cartel formation is not driven by reverse causality or omitted variable bias. Consequently, the results support H1: leverage positively predicts product cartel formation once firm-level characteristics are controlled for, consistent with financially constrained firms using collusion as a mechanism for revenue stabilization.

Given the change in the sign of leverage following the inclusion of control variables, this opens the possibility of segmenting the variable of interest into cumulative increases and cumulative decreases over time. The aim is to assess whether the sign reversal is driven, in

Table 9: IV Estimates – Product Market Cartel Formation and Leverage

	Aggregate leverage shift		Aggregate shift + Sales share	
	(1)	(2)	(3)	(4)
Leverage$_{t-1}$	-0.0171^{**}	0.0263^{***}	-0.0169^{**}	0.0260^{***}
	(0.0068)	(0.0101)	(0.0067)	(0.0100)
ROA$_{t-1}$		0.0001^{**}		0.0001^{**}
		(0.0000)		(0.0000)
ln(Sales)$_{t-1}$		0.0081^{***}		0.0080^{***}
		(0.0031)		(0.0030)
ln(Employees)$_{t-1}$		-0.0044^{**}		-0.0044^{**}
		(0.0018)		(0.0017)
HHI$_{t-1}$		0.0077		0.0077
		(0.0055)		(0.0054)
Year FE	✓	✓	✓	✓
Sector FE	✓	✓	✓	✓
Controls		✓		✓
KP LM p-value (underid)	0.000	0.000	–	–
KP Wald F (weak id)	26.51	22.42	17.29	15.55
SY critical value (10%)	16.38	16.38	19.93	19.93
AR p-value	0.015	–	0.039	–
Stock–Wright p-value	0.013	–	0.021	–
Hansen J p-value	–	–	0.072	0.110
Observations	4,371,271	3,875,211	4,371,271	3,875,211

Notes: Dependent variable is Cartel$_{it}^{PM}$. Aggregate leverage shift: the firm's lagged employment share within its two-digit CAE sector interacted with the economy-wide average leverage excluding the firm's own sector. Aggregate shift + Sales share: adds the firm's lagged sales share interacted with the leave-one-out industry average leverage as a second instrument, enabling the Hansen overidentification test. KP LM: Kleibergen–Paap rk LM statistic for underidentification. KP Wald F: Kleibergen–Paap weak identification test, cluster-robust. SY critical value: Stock–Yogo 10% maximal IV size threshold. AR: Anderson–Rubin test, robust to weak instruments. Stock–Wright S: weak-instrument-robust LM test. Hansen J: p-value of the overidentification test (H0: instruments are valid). Standard errors clustered at the firm level throughout. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

addition to omitted variable bias and the associated absence of control of observed heterogeneity at the firm level, by the failure to distinguish between increases and reductions in leverage. Table 10 examines whether the relationship between leverage and product cartel formation differs depending on the direction of leverage change, as specified in equation (3).

Under OLS without controls (column 1), the increasing-spell interaction is negative and significant, while the decreasing-spell coefficient is small and insignificant, reflecting the same omitted variable problem discussed above. Once controls are included (column 2), both interactions turn positive and significant, with the increasing-spell coefficient ($\hat{\beta}^{+} \approx 1.5 \times 10^{-5}$) exceeding the decreasing-spell coefficient ($\hat{\beta}^{-} \approx 1 \times 10^{-5}$). In terms of economic magnitude, a one-million-euro increase in leverage is associated with an increase of approximately 15 pp in the probability of collusion for firms on an upward leverage trajectory. The formal test of equality rejects the null hypothesis ($p = 0.027$), implying that the association between leverage and product collusion is stronger among firms that are accumulating debt than among those that are reducing it. This asymmetry is consistent with the financial stress mechanism: firms accumulating new liabilities face the sharpest incentive to stabilize revenues through product collusion, as servicing incremental debt requires predictable cash flows. This interpretation refines the baseline leverage effect in H1 by showing that it is not the level of leverage per se that drives collusive behavior, but rather the dynamics of debt accumulation: firms experiencing increases in leverage respond to heightened financial pressure by seeking more stable revenue streams, whereas reductions in leverage do not exhibit such a strong effect. Under Poisson, the pattern is less precise once controls are added.

Table 10: Product Market Cartel Formation and Leverage – Asymmetric Effects of Leverage Direction

	OLS		Poisson (PPML)	
	(1)	(2)	(3)	(4)
$\text{Leverage}_{i,t-1}$ × Increasing spell	-0.000004^{***}	0.000015^{***}	-0.2127^{***}	0.0437
	(0.000001)	(0.000002)	(0.0501)	(0.1014)
$\text{Leverage}_{i,t-1}$ × Decreasing spell	0.000001	0.000010^{***}	-0.0003	0.0249
	(0.000002)	(0.000003)	(0.0313)	(0.1778)
ROA_{t-1}		-0.000000		0.0008
		(0.000000)		(0.0009)
$\ln(\text{Sales})_{t-1}$		0.000071^{***}		1.1664^{***}
		(0.000014)		(0.0802)
$\ln(\text{Employees})_{t-1}$		0.000198^{***}		0.0074
		(0.000044)		(0.0808)
HHI_{t-1}		0.000449^{**}		6.1939^{***}
		(0.000223)		(1.1359)
Year FE	✓	✓	✓	✓
Sector FE	✓	✓	✓	✓
Controls		✓		✓
Test of asymmetry (OLS spec. 2): $F(1,\ 623{,}415) = 4.91,\quad p = 0.027$				
Observations	4,371,271	3,875,211	2,067,430	1,866,653

Notes: Dependent variable is Cartel_{it}^{PM}, a binary indicator equal to one when the firm is engaged in active product market collusion. $\text{Leverage}_{i,t-1}$ × Increasing spell is the product of lagged debt-to-assets and a dummy equal to one when the firm recorded a positive year-on-year change in leverage in the previous period. $\text{Leverage}_{i,t-1}$ × Decreasing spell is defined analogously for negative changes. Controls include ROA, log sales, log employees, and the Herfindahl–Hirschman index. All specifications include year and sector (CAE-2 digit) fixed effects. Standard errors clustered at the firm level in parentheses. Poisson estimates drop singleton observations separated by fixed effects. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Active product market cartel membership and leverage (H3). The analysis now reverses the causal direction. Table 11 presents the estimates of equation (7), which models leverage as the outcome variable during periods of active product market cartel participation. In the bivariate specification with sector and year fixed effects (column 1), the coefficient on product cartel participation is negative and highly significant ($\hat{\beta} = -0.545$, $p < 0.01$), suggesting that firms in active cartels carry lower leverage on average than peers in the same sector and year. This cross-sectional pattern, however, is driven by selection: firms that join cartels are systematically less leveraged than their peers to begin with, rather than reducing leverage as a consequence of cartel participation. Once firm fixed effects are introduced (columns 3–8), the coefficient reverses sign and becomes positive and significant across all specifications, ranging from 0.053 to 0.093. In the most demanding specification (column 8), the coefficient is 0.053, implying that within the same firm, active cartel participation is associated with an increase of approximately 0.053 percentage points in leverage relative to the mean. This finding *rejects* H3: Portuguese product market cartels do not use strategic deleveraging as a commitment device.

Table 11: Leverage During Ongoing Product Market Cartel – OLS with Multi-level Fixed Effects

	Sector + Year FE		Firm + Year FE		Firm + Year + Sector FE		Firm + Sector×Year FE	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Cartel_{it}^{PM}	-0.5450^{***}	1.2532^{***}	0.0908^{**}	0.0563^{*}	0.0902^{**}	0.0529^{*}	0.0926^{**}	0.0532^{*}
	(0.0458)	(0.0865)	(0.0414)	(0.0304)	(0.0419)	(0.0308)	(0.0363)	(0.0277)
ROA_{t-1}		-0.0024^{***}		-0.0006^{***}		-0.0006^{***}		-0.0006^{***}
		(0.0006)		(0.0001)		(0.0001)		(0.0001)
$\ln(\text{Sales})_{t-1}$		-0.4384^{***}		-0.1823^{***}		-0.1823^{***}		-0.1775^{***}
		(0.0029)		(0.0026)		(0.0026)		(0.0026)
$\ln(\text{Employees})_{t-1}$		0.2460^{***}		-0.0495^{***}		-0.0500^{***}		-0.0415^{***}
		(0.0037)		(0.0044)		(0.0044)		(0.0044)
HHI_{t-1}		-0.6806^{**}		0.8809^{***}		1.1776^{***}		-0.1089
		(0.2691)		(0.1454)		(0.1951)		(0.1694)
Sector FE	✓	✓			✓	✓		
Year FE	✓	✓	✓	✓	✓	✓		
Firm FE			✓	✓	✓	✓	✓	✓
Sector×Year FE							✓	✓
Controls		✓		✓		✓		✓
Observations	5,154,089	3,853,519	5,060,800	3,762,249	5,060,800	3,762,249	5,060,800	3,762,249

Notes: Dependent variable is Leverage_{it}, measured as debt-to-assets winsorized at the 1st and 99th percentiles. Cartel_{it}^{PM} is a binary indicator equal to one when the firm is engaged in active product market collusion. Controls include lagged ROA, log sales, log employees, and the Herfindahl–Hirschman index. Standard errors clustered at the firm level in parentheses. Singleton observations dropped where applicable. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table 12 replicates the same four fixed-effect structures using Poisson pseudo-maximum-likelihood estimation, which is more robust to the heavy right-skewness of leverage and handles zeros without requiring a log transformation. The pattern of results closely follows that of OLS across all specifications, with the sign reversal from negative to positive consistently emerging once firm fixed effects are introduced, reinforcing the robustness of the main finding.

Table 12: Leverage During Ongoing Product Market Cartel – Poisson (PPML) with Multilevel Fixed Effects

	Sector + Year FE		Firm + Year FE		Firm + Year + Sector FE		Firm + Sector×Year FE	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Cartel_{it}^{PM}	-0.5949^{***}	0.8327^{***}	0.0716^{*}	0.0503	0.0703^{*}	0.0470	0.0771^{**}	0.0454
	(0.0531)	(0.0790)	(0.0397)	(0.0311)	(0.0401)	(0.0317)	(0.0367)	(0.0291)
ROA_{t-1}		-0.0001^{***}		-0.0000^{***}		-0.0000^{***}		-0.0000^{***}
		(0.0000)		(0.0000)		(0.0000)		(0.0000)
$\ln(\text{Sales})_{t-1}$		-0.2760^{***}		-0.1113^{***}		-0.1112^{***}		-0.1074^{***}
		(0.0015)		(0.0015)		(0.0015)		(0.0015)
$\ln(\text{Employees})_{t-1}$		0.1025^{***}		-0.0315^{***}		-0.0316^{***}		-0.0249^{***}
		(0.0028)		(0.0031)		(0.0031)		(0.0031)
HHI_{t-1}		-0.4602^{**}		0.7591^{***}		0.8672^{***}		-0.1777
		(0.2107)		(0.1610)		(0.1930)		(0.2043)
Sector FE	✓	✓			✓	✓		
Year FE	✓	✓	✓	✓	✓	✓		
Firm FE			✓	✓	✓	✓	✓	✓
Sector×Year FE							✓	✓
Controls		✓		✓		✓		✓
Observations	5,154,089	3,853,519	5,052,439	3,759,983	5,052,439	3,759,983	5,052,439	3,759,983

Notes: Dependent variable is Leverage_{it}, measured as debt-to-assets winsorized at the 1st and 99th percentiles. Cartel_{it}^{PM} is a binary indicator equal to one when the firm is engaged in active product market collusion. Controls include lagged ROA, log sales, log employees, and the Herfindahl–Hirschman index. Singleton observations separated by fixed effects are dropped automatically by *ppmlhdfe*. Standard errors clustered at the firm level in parentheses. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Figure 1 presents the dynamic time fixed effects of product market cartel membership on leverage, estimated using the specification described in the methodology with firm and sector×year fixed effects (8). The coefficients capture the effect on leverage two and one year before the cartel onset ($t = -2$ and $t = -1$), at the year of onset ($t = 0$), and one and two years after ($t = +1$ and $t = +2$). At $t = -2$, the coefficient is positive and the confidence interval lies mostly above zero, suggesting that firms entering a cartel had somewhat higher leverage two years prior. By $t = -1$ and $t = 0$, the coefficients converge towards zero, consistent with leverage stabilizing in the immediate run-up to and at the moment of cartel formation. In the year following the onset ($t = +1$), the coefficient turns negative, which could reflect a short-term reduction in debt following the revenue stabilization that collusion provides, before reversing again at $t = +2$. The pattern that emerges is therefore one of elevated leverage in the pre-cartel period, a convergence towards zero around onset, and a non-monotonic adjustment thereafter. However, throughout the full window, the confidence intervals are wide and all coefficients are statistically indistinguishable from zero, meaning that the dynamic time fixed effects does not provide evidence of a systematic and precisely estimated dynamic pattern. The results are thus consistent with the static estimates, and the increase in leverage during active cartel spells remains the central finding for product market collusion, leading to a rejection of H3.

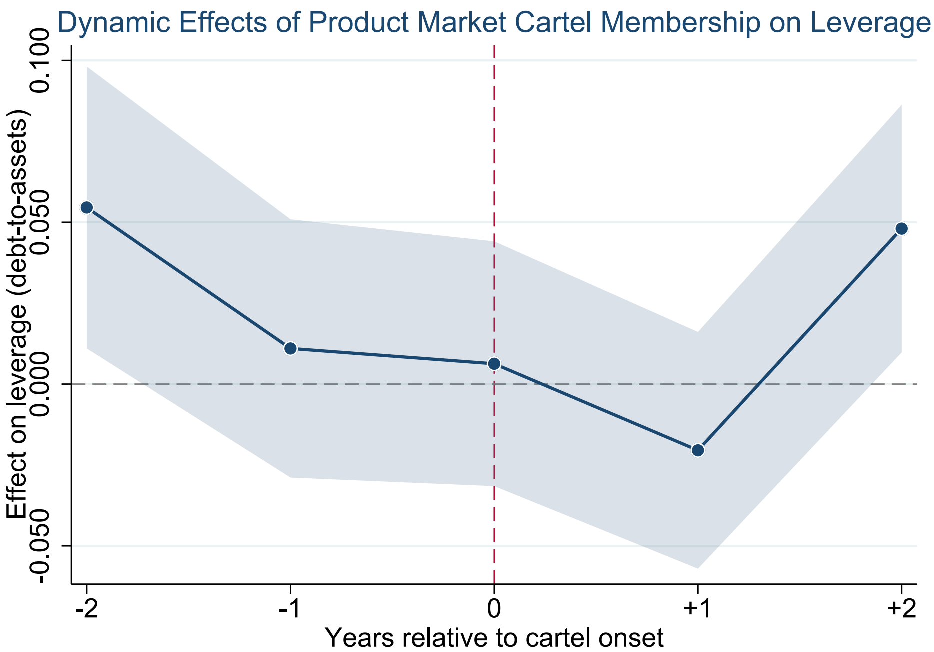

Figure 1: Dynamic Effects of Product Market Cartel Membership on Leverage

Notes: Coefficients from an OLS dynamic time fixed effects specification of leverage on leads and lags of product market cartel membership, with firm and sector×year fixed effects. Controls include lagged ROA, log sales, log employees, and the Herfindahl–Hirschman index. The shaded area represents 95% confidence intervals. Standard errors clustered at the firm level.

6.2 Labor Market Cartel

Leverage and labor cartel formation (H2). Table 13 reports the baseline estimates for the relationship between leverage and labor market cartel participation, identified from cases pursued by the AdC involving collusion in labor market, following equation (2). In the bivariate specification (columns 1 and 3), the coefficient of leverage is statistically insignificant under both OLS and Poisson, which again reflects the omitted variable problem discussed in the product market collusion analysis. Once controls are included (columns 2 and 4), the coefficient turns positive and significant under both estimators. Under OLS (column

2), a one-million-euro increase in leverage relative to the mean is associated with an increase of approximately 7 percentage points in the probability of labor market cartel participation ($\hat{\beta} \approx 7 \times 10^{-6}$, $p < 0.01$). The Poisson specification follows the same pattern, with the coefficient turning positive and significant once controls are added ($\hat{\beta} \approx 0.183$, $p < 0.01$). These results clearly support H2, thereby confirming that the positive leverage–cartel formation relationship is not specific to the product market.

Table 13: Labor Market Cartel Formation and Leverage

	OLS		Poisson (PPML)	
	(1)	(2)	(3)	(4)
Leverage$_{t-1}$	-0.000001	0.000007***	-0.0226	0.1830***
	(0.000001)	(0.000002)	(0.0385)	(0.0358)
ROA$_{t-1}$		0.000000**		0.0019*
		(0.000000)		(0.0011)
ln(Sales)$_{t-1}$		0.000003		0.3453
		(0.000005)		(0.2572)
ln(Employees)$_{t-1}$		0.000118***		1.0845***
		(0.000030)		(0.3113)
HHI$_{t-1}$		0.000221		-6.7236
		(0.000165)		(5.8277)
Year FE	✓	✓	✓	✓
Sector FE	✓	✓	✓	✓
Controls		✓		✓
Observations	4,371,271	3,875,211	497,869	441,365

Notes: Dependent variable is Cartel$_{it}^{LM}$, a binary indicator equal to one when the firm is engaged in active labor market collusion. All specifications include year and sector (CAE-2 digit) fixed effects. Leverage$_{i,t-1}$ is measured by lagged debt-to-assets, winsorized at the 1st and 99th percentiles. Controls include ROA, log sales, log employees, and the Herfindahl–Hirschman index. Standard errors clustered at the firm level in parentheses. Poisson estimates drop singleton observations separated by fixed effects. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

The endogeneity test reported in Table 21 rejects the exogeneity of leverage in the labor market cartel formation, motivating the Bartik instrument described in equations (4), (5), and (6).

Table 14 reports the IV estimates. The Kleibergen–Paap LM and Wald F statistics indicate that the instrument is relevant and sufficiently strong. The Anderson–Rubin test rejects the null of no leverage effect on labor market cartel formation, while the Hansen J test does not reject instrument validity. However, the endogeneity test fails to reject exogeneity in the overidentified case (Table 21, column (4)). The second-stage leverage coefficient is positive and significant once controls are included, consistent with the OLS results. Although smaller than in product market cartel analysis, the positive relationship between leverage and labor market cartel formation is unlikely to be driven by reverse causality or omitted variables.

Table 14: IV Estimates – Labor Market Cartel Formation and Leverage

	Aggregate leverage shift		Aggregate shift + Sales share	
	(1)	(2)	(3)	(4)
Leverage$_{t-1}$	-0.0027^{***}	0.0033^{**}	-0.0027^{***}	0.0032^{**}
	(0.0010)	(0.0015)	(0.0010)	(0.0015)
ROA$_{t-1}$		0.0000^{*}		0.0000^{*}
		(0.0000)		(0.0000)
ln(Sales)$_{t-1}$		0.0010^{**}		0.0010^{**}
		(0.0005)		(0.0004)
ln(Employees)$_{t-1}$		-0.0005^{*}		-0.0004^{*}
		(0.0002)		(0.0002)
HHI$_{t-1}$		0.0006		0.0006
		(0.0006)		(0.0006)
Year FE	✓	✓	✓	✓
Sector FE	✓	✓	✓	✓
Controls		✓		✓
KP LM p-value (underid)	0.000	0.000	–	–
KP Wald F (weak id)	26.51	22.42	17.29	15.55
SY critical value (10%)	16.38	16.38	19.93	19.93
AR p-value	0.023	–	0.059	–
Stock–Wright p-value	0.007	–	0.019	–
Hansen J p-value	–	–	0.143	0.251
Observations	4,371,271	3,875,211	4,371,271	3,875,211

Notes: The dependent variable is Cartel$_{it}^{LM}$. The main instrument is a firm's lagged employment share within its two-digit CAE sector interacted with the economy-wide average leverage excluding that sector. A second specification adds an analogous sales-share instrument, allowing the Hansen overidentification test (H0: instruments are valid). Reported diagnostics include the Kleibergen–Paap LM statistic (underidentification), Kleibergen– Paap Wald F statistic (weak identification), the Stock–Yogo critical value, the Anderson–Rubin and Stock–Wright weak-instrument-robust tests, and the Hansen J test. Significance levels: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table 15 replicates the asymmetric decomposition of equation (3) for labor market cartel formation. In contrast to the product market cartel case, the decreasing-spell coefficient is larger than the increasing-spell coefficient in OLS ($\hat{\beta}^{-} \approx 1.1 \times 10^{-5}$ versus $\hat{\beta}^{+} \approx 5 \times 10^{-6}$). In terms of economic magnitude, a one-million-euro increase in leverage relative to the mean is associated with an increase of approximately 11 p.p. in the probability of labor market cartel participation for firms on a downward leverage trajectory. This pattern contrasts with the product market cartel results and suggests that for labor market cartels it may be firms that are already reducing their debt load, and therefore facing tighter margins, that have stronger incentives to coordinate on labor costs. The formal test of equality does not reject the null hypothesis ($p = 0.151$), suggesting that leverage affects labor cartel participation similarly regardless of whether firms are accumulating or reducing debt. Under Poisson, the pattern is broadly consistent with OLS.

Table 15: Labor Market Cartel Formation and Leverage – Asymmetric Effects of Leverage Direction

	OLS		Poisson (PPML)	
	(1)	(2)	(3)	(4)
Leverage$_{i,t-1}$ × Increasing spell	-0.000001	0.000005***	-0.0333	0.1539***
	(0.000001)	(0.000002)	(0.0414)	(0.0415)
Leverage$_{i,t-1}$ × Decreasing spell	0.000007*	0.000011**	0.0946***	0.2857***
	(0.000004)	(0.000005)	(0.0289)	(0.0392)
ROA$_{t-1}$		0.000000**		0.0031***
		(0.000000)		(0.0010)
ln(Sales)$_{t-1}$		0.000002		0.3419
		(0.000005)		(0.2590)
ln(Employees)$_{t-1}$		0.000119***		1.0896***
		(0.000030)		(0.3129)
HHI$_{t-1}$		0.000218		-6.9319
		(0.000165)		(5.8647)
Year FE	✓	✓	✓	✓
Sector FE	✓	✓	✓	✓
Controls		✓		✓
Test of asymmetry (OLS spec. 2):	$F(1,\ 623{,}415) = 2.06,\quad p = 0.151$			
Observations	4,371,271	3,875,211	497,869	441,365

Notes: Dependent variable is Cartel$^{LM}_{it}$, a binary indicator equal to one when the firm is engaged in active labor market collusion. Leverage$_{i,t-1}$ × Increasing spell is the product of lagged debt-to-assets and a dummy equal to one when the firm recorded a positive year-on-year change in leverage in the previous period. Leverage$_{i,t-1}$ × Decreasing spell is defined analogously for negative changes. All specifications include year and sector fixed effects. Standard errors clustered. Poisson estimates drop singleton observations separated by fixed effects. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Active labor market cartel membership and leverage (H4). Table 16 replicates the multi-level fixed-effect analysis of equation (7), following the same four fixed-effect structures described above. The results diverge markedly from the product market cartel case. The coefficient on labor market cartel participation is negative and statistically significant across all specifications that exploit within-firm variation, with the exception of column 2, where the coefficient is positive (0.917). Once firm fixed effects are introduced and identification relies on within-firm variation, the coefficient turns negative and remains so across all remaining specifications. In the most demanding specification (column 8), which absorbs firm fixed effects and sector×year fixed effects, the coefficient is -0.756 ($p < 0.01$), implying that within the same firm, leverage, relative to the mean, is approximately 0.756 percentage points lower during labor market cartel periods than in non-cartel years, after controlling for all sector-specific annual shocks. This result *supports* H4 and yet contrasts directly with the product market cartel findings. This is perhaps surprising given that Gonzaga et al. (2013) show that labor market collusion produces outcomes equivalent to product market collusion, with higher prices. The divergence in results suggests that the mechanism linking debt and cartel participation operates differently across the two types of cartel, consistent with the Maksimovic (1988) framework applying asymmetrically: product cartels generate abnormal revenues that can service debt, whereas labor cartels generate cost savings whose implications for leverage are less direct.

Table 16: Leverage During Ongoing Labor Market Cartel – OLS with Multi-level Fixed Effects

	Sector + Year FE		Firm + Year FE		Firm + Year + Sector FE		Firm + Sector×Year FE	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Cartel_{it}^{LM}	-0.5825^{***}	0.9172^{***}	-0.6463^{***}	-0.5472^{**}	-0.6529^{***}	-0.5545^{**}	-0.9480^{***}	-0.7556^{***}
	(0.1671)	(0.1512)	(0.2421)	(0.2569)	(0.2421)	(0.2568)	(0.2615)	(0.2699)
ROA_{t-1}		-0.0024^{***}		-0.0006^{***}		-0.0006^{***}		-0.0006^{***}
		(0.0006)		(0.0001)		(0.0001)		(0.0001)
$\ln(\text{Sales})_{t-1}$		-0.4383^{***}		-0.1823^{***}		-0.1823^{***}		-0.1775^{***}
		(0.0029)		(0.0026)		(0.0026)		(0.0026)
$\ln(\text{Employees})_{t-1}$		0.2462^{***}		-0.0495^{***}		-0.0500^{***}		-0.0415^{***}
		(0.0037)		(0.0044)		(0.0044)		(0.0044)
HHI_{t-1}		-0.6752^{**}		0.8813^{***}		1.1780^{***}		-0.1087
		(0.2691)		(0.1454)		(0.1951)		(0.1694)
Sector FE	✓	✓			✓	✓		
Year FE	✓	✓	✓	✓	✓	✓		
Firm FE			✓	✓	✓	✓	✓	✓
Sector×Year FE							✓	✓
Controls		✓		✓		✓		✓
Observations	5,154,089	3,853,519	5,060,800	3,762,249	5,060,800	3,762,249	5,060,800	3,762,249

Notes: Dependent variable is Leverage_{it}, measured as debt-to-assets winsorized at the 1st and 99th percentiles. Cartel_{it}^{LM} is a binary indicator equal to one when the firm is engaged in active labor market collusion. Controls include lagged ROA, log sales, log employees, and the Herfindahl–Hirschman index. Standard errors clustered at the firm level in parentheses. Singleton observations dropped where applicable. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table 17 replicates the analysis using Poisson pseudo-maximum-likelihood estimation. The results confirm the OLS estimation: the coefficient on labor market collusion remains negative and statistically significant across all within-firm specifications, while column 2 again yields a positive coefficient (0.711). The within-firm estimates range from -0.327 to -0.534, somewhat smaller in magnitude than the OLS but qualitatively identical in sign and significance.

Table 17: Leverage During Ongoing Labor Market Cartel – Poisson (PPML) with Multi-level Fixed Effects

	Sector + Year FE		Firm + Year FE		Firm + Year + Sector FE		Firm + Sector×Year FE	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Cartel_{it}^{LM}	-0.3798^{***}	0.7109^{***}	-0.4082^{***}	-0.3266^{***}	-0.4099^{***}	-0.3288^{***}	-0.5343^{***}	-0.4356^{***}
	(0.1316)	(0.0936)	(0.1349)	(0.1184)	(0.1351)	(0.1186)	(0.1362)	(0.1200)
ROA_{t-1}		-0.0001^{***}		-0.0000^{***}		-0.0000^{***}		-0.0000^{***}
		(0.0000)		(0.0000)		(0.0000)		(0.0000)
$\ln(\text{Sales})_{t-1}$		-0.2760^{***}		-0.1113^{***}		-0.1112^{***}		-0.1074^{***}
		(0.0015)		(0.0015)		(0.0015)		(0.0015)
$\ln(\text{Employees})_{t-1}$		0.1025^{***}		-0.0315^{***}		-0.0316^{***}		-0.0249^{***}
		(0.0028)		(0.0031)		(0.0031)		(0.0031)
HHI_{t-1}		-0.4597^{**}		0.7596^{***}		0.8677^{***}		-0.1776
		(0.2107)		(0.1610)		(0.1930)		(0.2043)
Sector FE	✓	✓			✓	✓		
Year FE	✓	✓	✓	✓	✓	✓		
Firm FE			✓	✓	✓	✓	✓	✓
Sector×Year FE							✓	✓
Controls		✓		✓		✓		✓
Observations	5,154,089	3,853,519	5,052,439	3,759,983	5,052,439	3,759,983	5,052,439	3,759,983

Notes: Dependent variable is Leverage_{it}, measured as debt-to-assets winsorized at the 1st and 99th percentiles. Cartel_{it}^{LM} is a binary indicator equal to one when the firm is engaged in active labor market collusion. Controls include lagged ROA, log sales, log employees, and the Herfindahl–Hirschman index. Singleton observations separated by fixed effects are dropped automatically by *ppmlhdfe*. Standard errors clustered at the firm level in parentheses. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Figure 2 presents the dynamic time fixed effects of labor market cartel membership on leverage, estimated using the same specification as Figure 1. The pattern differs markedly from the product market collusion case. In the two years prior to cartel onset, the coefficients are close to zero or slightly positive, and a joint linear combination test of the pre-entry and post-entry coefficients confirms that the two sets differ significantly, indicating that the leverage dynamics documented after cartel onset reflect the effect of cartel participation rather than a pre-existing trajectory. At $t = 0$, the coefficient drops sharply to approximately -0.8, indicating a pronounced reduction in leverage in the year the cartel begins. In the years following onset, the coefficient recovers towards zero but remains negative. The absence of pre-entry dynamics and the sharp drop at $t = 0$ suggest that leverage falls because the cartel is formed, consistent with strategic deleveraging to sustain the labor cartel agreement. However, the wide confidence intervals throughout reflect the small number of AdC labor market episodes in the sample, and no individual coefficient is statistically significant at conventional levels, implying that this interpretation should be treated with caution.

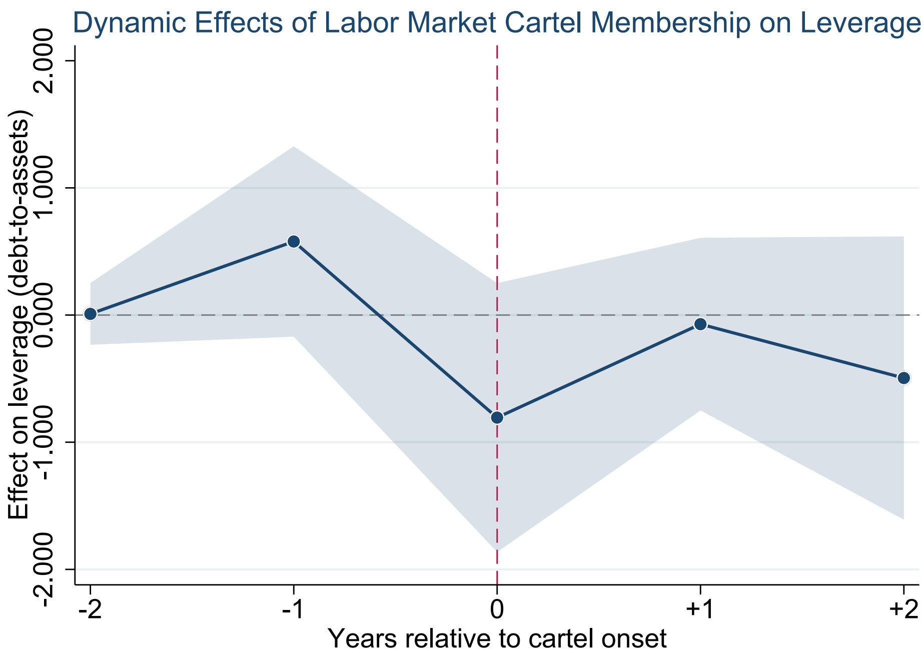

Figure 2: Dynamic Effects of Labor Market Cartel Membership on Leverage

Notes: Coefficients from an OLS dynamic time fixed effects specification of leverage on leads and lags of labor market cartel membership, with firm and sector×year fixed effects. Controls include lagged ROA, log sales, log employees, and the Herfindahl–Hirschman index. The shaded area represents 95% confidence intervals. Standard errors clustered at the firm level.

6.3 Minimum Wage Collusion

Unlike the AdC-based measures used above, collusion is now identified through the use of the statutory minimum wage as a tacit focal point (Shelkova, 2015).

Table 18 presents the OLS estimates of equation (10). Without the imposition of firm fixed effects, the coefficient is positive and significant: firms with at least one affected worker carry leverage relative to the mean approximately 9 pp higher than peers in the same sector and year. Once firm fixed effects are introduced, the coefficient remains positive but loses statistical significance across all within-firm specifications, suggesting that the association is driven primarily by persistent between-firm differences rather than within-firm changes. Unlike the case of AdC labor market collusion, where the sign remained consistently negative within firms, the positive sign here is preserved, suggesting that firms engaged in tacit wage coordination through the minimum wage tend to carry higher leverage, which is inconsistent with H5. The contrast may reflect differences in the nature of the two measures: the AdC indicator captures formally prosecuted agreements, whereas the minimum wage collusion indicator is a broader, indirect measure of tacit coordination covering a much larger share of firms.

Table 18: Leverage During Ongoing Minimum Wage Collusion – OLS with Multi-level Fixed Effects

	Sector + Year FE		Firm + Year FE		Firm + Year + Sector FE		Firm + Sector×Year FE	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\text{Collusion}_{it}^{MW}$	0.0899***	0.0575***	0.0033*	0.0027	0.0033*	0.0026	0.0031	0.0021
	(0.0032)	(0.0034)	(0.0020)	(0.0020)	(0.0020)	(0.0020)	(0.0020)	(0.0020)
ROA_{t-1}		−0.0008*		−0.0003***		−0.0003***		−0.0003***
		(0.0004)		(0.0001)		(0.0001)		(0.0001)
$\ln(\text{Sales})_{t-1}$		−0.2930***		−0.1550***		−0.1550***		−0.1519***
		(0.0026)		(0.0023)		(0.0023)		(0.0023)
$\ln(\text{Employees})_{t-1}$		0.1779***		0.0350***		0.0349***		0.0379***
		(0.0030)		(0.0028)		(0.0028)		(0.0028)
HHI_{t-1}		−0.2468**		0.0913		0.1389		−0.0090
		(0.1124)		(0.0679)		(0.0883)		(0.0634)
Sector FE	✓	✓			✓	✓		
Year FE	✓	✓	✓	✓	✓	✓		
Firm FE			✓	✓	✓	✓	✓	✓
Sector×Year FE							✓	✓
Controls		✓		✓		✓		✓
Observations	2,533,619	2,013,399	2,464,976	1,959,080	2,464,976	1,959,080	2,464,976	1,959,080

Notes: Dependent variable is Leverage_{it}, measured as debt-to-assets winsorized at the 1st and 99th percentiles. $\text{Collusion}_{it}^{MW}$ is a binary indicator equal to one if the firm has at least one worker whose observed wage equals the statutory minimum but whose predicted latent wage exceeds it (Shelkova, 2015). The latent wage equation follows equation (9). Controls include lagged ROA, log sales, log employees, and the Herfindahl–Hirschman index. Standard errors clustered at the firm level in parentheses. Singleton observations dropped where applicable. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table 19 replicates the analysis using Poisson pseudo-maximum-likelihood estimation. The results are more robust than the OLS findings: the coefficient on minimum wage collusion is positive and statistically significant across all eight specifications, suggesting that the loss of significance observed under OLS reflects sensitivity to extreme observations rather than a genuine absence of effect.

Table 20 examines the sensitivity of the baseline result to three alternative operationalizations of the collusion indicator. Column 1 replaces the binary indicator with the continuous share of affected workers in total firm employment, and the coefficient is positive and highly significant, indicating that the association between minimum wage collusion and leverage strengthens as the fraction of the workforce subject to wage suppression increases. Column 2 uses a binary indicator requiring at least ten percent of the firm's workers to satisfy the collusion criterion and yields a positive but marginally insignificant coefficient. Column 3 uses a binary indicator based on whether the share of affected workers exceeds the sample median among firms with at least one affected worker, yielding a large and significant coeffi-

Table 19: Leverage During Ongoing Minimum Wage Collusion – Poisson (PPML) with Multi-level Fixed Effects

	Sector + Year FE		Firm + Year FE		Firm + Year + Sector FE		Firm + Sector×Year FE	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\text{Collusion}_{it}^{MW}$	0.0915***	0.0732***	0.0041**	0.0047**	0.0041**	0.0046**	0.0051**	0.0046**
	(0.0032)	(0.0035)	(0.0021)	(0.0022)	(0.0021)	(0.0022)	(0.0020)	(0.0021)
ROA_{t-1}		−0.0001***		−0.0000***		−0.0000***		−0.0000***
		(0.0000)		(0.0000)		(0.0000)		(0.0000)
$\ln(\text{Sales})_{t-1}$		−0.2925***		−0.1440***		−0.1439***		−0.1382***
		(0.0025)		(0.0020)		(0.0020)		(0.0020)
$\ln(\text{Employees})_{t-1}$		0.1494***		0.0289***		0.0289***		0.0327***
		(0.0034)		(0.0028)		(0.0028)		(0.0028)
HHI_{t-1}		−0.2355*		0.1391		0.2158*		−0.0113
		(0.1364)		(0.0971)		(0.1175)		(0.0926)
Sector FE	✓	✓			✓	✓		
Year FE	✓	✓	✓	✓	✓	✓		
Firm FE			✓	✓	✓	✓	✓	✓
Sector×Year FE							✓	✓
Controls		✓		✓		✓		✓
Observations	2,533,619	2,013,399	2,464,976	1,959,080	2,464,976	1,959,080	2,464,976	1,959,080

Notes: Dependent variable is Leverage_{it}, measured as debt-to-assets winsorized at the 1st and 99th percentiles. $\text{Collusion}_{it}^{MW}$ is a binary indicator equal to one if the firm has at least one worker whose observed wage equals the statutory minimum but whose predicted latent wage exceeds it (Shelkova, 2015). Controls include lagged ROA, log sales, log employees, and the Herfindahl–Hirschman index. Singleton observations separated by fixed effects are dropped automatically by *ppmlhdfe*. Standard errors clustered at the firm level in parentheses. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

cient. Across all three alternative measures, the direction of the effect is positive, providing consistent evidence that minimum wage collusion is associated with higher leverage and confirming that the baseline result is not an artifact of the chosen operationalization, although the association remains inconsistent with H5.

Figure 3 presents the dynamic time fixed effects for minimum wage collusion, using the same specification (8). The pre-period coefficients at $t = -2$ and $t = -1$ are positive and statistically significant, including the pre-collusion periods. This persistent behavior weakens the interpretation of the estimated relationship as arising from collusion itself: firms classified as minimum wage colluders already exhibited higher leverage two years before the indicator turns on. Rather than capturing a discrete shock with a clean onset, the minimum wage collusion indicator reflects a persistent firm characteristic that is more strongly expressed in the pre-period than after: firms that systematically pay workers at the minimum despite their productive characteristics tend to carry higher leverage throughout the observation window. The associations reported in Tables 18–20 should therefore be interpreted as descriptive correlations.

Table 20: Robustness – Alternative Minimum Wage Collusion Measures

	Continuous (1)	Binary (10%) (2)	Binary (median) (3)
Share of affected workers	0.0374***		
	(0.0038)		
$\text{Collusion}_{it}^{MW}$ ($\geq 10\%$)		0.0033	
		(0.0020)	
$\text{Collusion}_{it}^{MW}$ (above median)			0.0308***
			(0.0031)
ROA_{t-1}	−0.0003***	−0.0003***	−0.0003***
	(0.0001)	(0.0001)	(0.0001)
$\ln(\text{Sales})_{t-1}$	−0.1512***	−0.1519***	−0.1513***
	(0.0023)	(0.0023)	(0.0023)
$\ln(\text{Employees})_{t-1}$	0.0388***	0.0380***	0.0390***
	(0.0028)	(0.0028)	(0.0028)
HHI_{t-1}	−0.0093	−0.0090	−0.0095
	(0.0634)	(0.0634)	(0.0633)
Firm FE	✓	✓	✓
Sector×Year FE	✓	✓	✓
Observations	1,959,080	1,959,080	1,959,080

Notes: Dependent variable is Leverage_{it}. Continuous: share of workers with observed wage equal to the minimum and predicted latent wage above the minimum in total firm employment. Binary (10%): share exceeds 10%. Binary (median): share exceeds the sample median among firms with at least one affected worker. Standard errors clustered at the firm level in parentheses. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

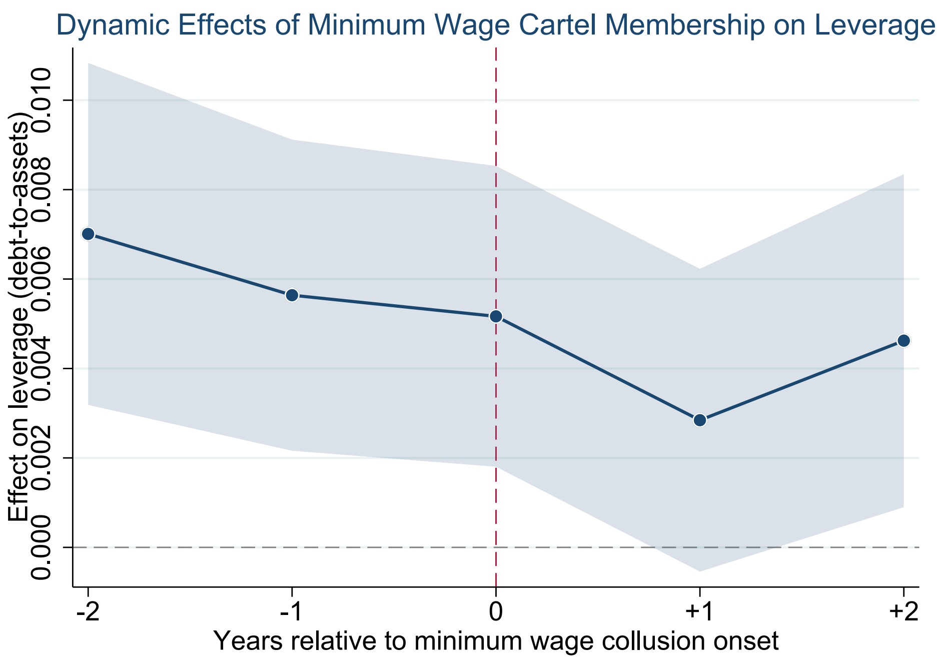

Figure 3: Dynamic Effects of Minimum Wage Collusion on Leverage

Notes: Coefficients from an OLS dynamic time fixed effects specification of leverage on leads and lags of the minimum wage collusion, with firm and sector×year fixed effects. Controls include lagged ROA, log sales, log employees, and the HHI. The shaded area represents 95% confidence intervals. Standard errors clustered at the firm level.

7 Discussion

7.1 Scholarly Implications

The findings contribute to, and depart from, several strands of prior work in ways that refine theoretical understanding of the leverage–collusion nexus. The original prediction of Maksimovic (1988) is that more leveraged firms are *less* likely to form cartels because debt erodes the credibility of punishment threats. Prior empirical work offered mixed but broadly supportive evidence: Levenstein and Suslow (2016) find a negative association between industry leverage and US cartel formation, and Ferrés et al. (2021) document within-firm deleveraging during active international cartel spells.

The results on the impact of pre-entry firm leverage on cartel formation challenge these findings, at least conditionally. Regardless of the cartel type, any negative leverage–formation relationship disappears and reverses once firm-level controls are added to the MLRM's deterministic component, suggesting that the raw negative correlation in prior work may partially reflect firm-size selection: large firms are less leveraged on average and more likely to operate in concentrated markets where cartels form. Once firm size and market concentration are controlled, leverage enters positively and survives Bartik instrumentation, pointing to the financial stress mechanism articulated by Huric-Larsen (2025): firms accumulating new debt seek collusion to stabilize the revenue streams needed to service that debt. This reinterpretation shifts the empirical consensus from the participation constraint to the strategic-demand-for-collusion channel, a distinction with important implications for theoretical models of cartel entry.

Furthermore, the result most clearly at odds with the prior literature is the *rising* leverage during active product cartel spells. Ferrés et al. (2021), using the PIC database, find that cartel firms reduce book leverage by approximately 3 pp, driven by debt repayment and concentrated in the first two years of collusion. The opposite is found for Portuguese product cartels. Several interpretations reconcile the discrepancy. First, the international cartel database is dominated by large-firm, global cartels whose members have the financial flexibility to retire debt strategically, while Portuguese product cartels include a substantial share of small and medium enterprises (SMEs) in services and domestic markets, for whom collusion generates revenues that are absorbed into working capital or investment rather than debt repayment. Second, Forsbacka et al. (2025) show that macroeconomic conditions shape cartel dynamics: the sample covers the post-crisis austerity period (2010–2015) and the COVID shock (2020–2022), both of which may have created incentives to increase leverage precisely during cartel spells to fund operations when external financing was tight. Third, and most importantly, the findings for labor cartels—which do show strategic deleveraging—suggest that the Maksimovic (1988) mechanism is not absent from the Portuguese institutional context but is specific to the cartel type. This is a novel theoretical insight: the direction of financial adjustment during collusion depends on whether coordination occurs on the output or the input side of the market.

To the best of my knowledge, this is the first study to examine the leverage consequences of labor market cartels, extending a literature that has focused exclusively on product market coordination. Delabastita and Rubens (2025) show that employer collusion suppresses wages below the competitive benchmark, and Gonzaga et al. (2013) establish the structural equivalence between product and labor market collusion. The results confirm that the leverage implications of these two forms of collusion are not symmetric. The cash flow channel differs fundamentally: product cartels generate revenue gains that may be reinvested, while labor cartels generate cost savings that flow more directly to free cash flow available for debt repayment. This distinction suggests that future theoretical models of collusion and financial structure should incorporate the input-output asymmetry explicitly. The labor-cartel deleveraging result is also consistent with the prediction of Brander and Lewis (1986) that competitive commitment requires low leverage, but it reverses the sign for product markets where supra-competitive revenues can dominate the commitment motive.

The minimum wage collusion analysis extends Shelkova (2015)'s focal-point framework to the Portuguese context and establishes, for the first time, a connection between this form of tacit coordination and corporate financial structure. While the pre-trend coefficients are positive and significant two years before the indicator turns on, undermining a causal interpretation, the persistent positive association between minimum wage dependence and leverage is informative as a stylized fact: firms that systematically pay at the statutory floor despite their workers' productive characteristics are also persistently more leveraged than comparable firms. This co-occurrence is consistent with a firm-type hypothesis—small, credit-constrained, labor-intensive firms that compete primarily on cost—rather than a causal chain from coordination to leverage. Future research exploiting regional or sector-specific minimum wage variation, or the natural experiments generated by large discrete minimum wage increases, could provide a sharper identification to establish causality. The Portuguese setting is particularly well-suited for this, given the series of substantial nominal minimum wage increases over the sample period.

7.2 Managerial Implications

The results carry several implications for corporate governance, financial management, and strategic decision-making. First, the positive association between pre-entry firm leverage and cartel formation implies that capital structure choices interact with competitive behavior in ways that boards and chief financial officers should internalize. A firm that increases its debt load to fund expansion or acquisitions may be simultaneously increasing the internal pressure on management to seek revenue stabilization through anti-competitive coordination. Governance structures should therefore integrate competitive compliance risk into the assessment of leverage policy, particularly in concentrated industries where collusion is structurally more tempting. The asymmetric result—that rising-leverage firms are more strongly associated with cartel formation than firms reducing debt—suggests that compliance monitoring should be especially attentive to periods of rapid debt accumulation.

Second, the finding that labor cartel firms strategically deleverage has a practical corollary: a firm that is actively participating in a no-poach or wage-fixing agreement may exhibit an anomalous reduction in leverage relative to sector peers, as it repays debt using wage-bill savings. Conversely, firms with persistently high leverage and minimum-wage dependence—as identified by the focal-point indicator—represent a combined risk profile for compliance officers. Internal audit and compliance functions could incorporate leverage trajectory analysis into labor-market conduct monitoring, particularly in sectors with high minimum-wage incidence, such as retail, hospitality, and food manufacturing.

Third, the evidence also suggests that leverage in concentrated industries is not purely a reflection of financing costs and asset tangibility. Instead, it also encodes competitive conduct. Credit analysts and lending institutions assessing firms in oligopolistic markets should be aware that a firm's leverage trajectory may be shaped by collusive coordination. During active product cartel spells, leverage may rise as collusive revenues are reinvested rather than used for debt reduction. The improved cash flow from supra-competitive rents services the increased debt load, masking the underlying competitive conduct. During labor cartel spells, the opposite may occur: anomalous deleveraging in the absence of observable revenue growth or asset sales could indicate that cost savings from wage coordination are being directed to debt repayment. These patterns could in principle be incorporated into credit monitoring systems in concentrated industries.

Lastly, the financial stress mechanism—where leveraged firms seek collusion to stabilize revenues—has specific relevance for leveraged buyouts and post-acquisition integration. Private equity transactions typically involve sharp increases in leverage, and the evidence suggests that such increases raise the probability of cartel participation conditional on observable firm characteristics. Acquirers and private equity sponsors should strengthen antitrust compliance frameworks in proportion to the leverage increase imposed by the transaction under scrutiny, particularly when the acquired firm operates in concentrated markets.

7.3 Policy Implications

A natural implication of the Maksimovic (1988) framework, which has arguably influenced enforcement practice, is that highly leveraged firms are poor cartel candidates because they cannot credibly commit to sustaining long-horizon punishment threats. The evidence presented here provides strong empirical grounds to reject this heuristic. In both product and labor markets, higher leverage is positively associated with cartel formation once firm size and market concentration are controlled, and this relationship survives instrumentation. Competition authorities—including the AdC, the European Commission, and national agencies conducting ex-ante market monitoring—should not treat high leverage as an indicator of reduced collusion risk. If anything, financially stressed firms in concentrated markets represent an elevated collusion risk, and leverage accumulation in such markets should attract attention rather than reassurance.

Moreover, the findings suggest that financial data, which competition authorities can access through compulsory disclosure requirements or cooperation with financial supervisors, could usefully complement market structure indicators in risk-based cartel screening. Firms exhibiting simultaneously high leverage, rapid debt accumulation, operating in concentrated sectors, and lacking significant export diversification may warrant earlier investigative attention than the current enforcement practice of responding primarily to leniency applications. The shift-share instrument used in this study confirms a tractable approach to isolating exogenous variation in leverage from aggregate financial conditions, which could be incorporated into algorithmic screening tools.

The AdC's two labor market decisions—the first such enforcement actions in Portugal—had measurable and statistically significant financial consequences for the sanctioned firms. Active labor cartel membership is associated with a reduction of approximately 0.76 pp in leverage within the same firm, and the dynamic time fixed effects shows that this effect originates at cartel onset rather than pre-existing it. This suggests that labor market enforcement has effects beyond the direct wage consequences: it disrupts the financial adjustment strategy through which labor cartel members sustain their agreements. Jurisdictions considering whether to prioritize labor market enforcement—including the European Commission, which has recently signaled interest in no-poach agreements—should anticipate that enforcement will have ripple effects on the financial structure of sanctioned firms that may reinforce the deterrent effect by eliminating the cost savings used to service or retire debt.

Additionally, the persistent correlation between minimum wage dependence and leverage has implications for the design of minimum wage policy. Firms that systematically pay the statutory floor and carry above-average leverage represent a financially constrained, cost-sensitive segment of the economy. Large discrete minimum wage increases—such as those observed in Portugal between 2017 and 2019—may interact with leverage in complex ways: on one hand, they compress margins for constrained firms, potentially amplifying financial distress. On the other hand, the evidence of Liang (2024) and Arabzadeh et al. (2024) suggests that minimum wage increases cause firms to reduce leverage by limiting the wage-bill flexibility that underlies debt-service capacity. Governmental decisions designing the trajectory of minimum wage increases should consider these financial structure effects alongside the more commonly studied employment and wage distribution effects, particularly for highly leveraged firms in low-wage sectors.

Ultimately, the findings highlight a natural complementarity between information held by banking supervisors—who have access to firm-level leverage data through credit registers and supervisory reporting—and competition authorities that rely primarily on market structure data and leniency applications. In the European context, the Single Supervisory Mechanism (SSM) under the European Central Bank maintains granular firm-level financial data that, subject to appropriate legal and confidentiality frameworks, could support risk-based cartel screening by national competition authorities. The positive leverage–cartel formation relationship documented here provides an empirical justification for developing such coordination mechanisms, particularly for critical sectors of the economic activity.

7.4 Limitations

Several limitations remain, and several directions for future research can be identified. First, the AdC's enforcement record contains only two labor market decisions over the 2010–2022 period, covering 35 unique firms and 114 firm-year observations. While the consistency of results across many fixed-effect specifications and two methods (OLS and PPML) provides some reassurance, the confidence intervals in the dynamic time fixed effects for labor cartel membership are wide, and no individual coefficient achieves conventional statistical significance. The effect magnitudes are economically large (approximately 0.76 pp in the most demanding specification), but the precision of the estimate depends on a small number of enforcement episodes. Future work may aggregate labor market enforcement records across multiple jurisdictions to generate a pooled sample with sufficient power for precise inference.

Second, the minimum wage focal-point strategy was applied as a proxy for tacit wage coordination, but the indicator cannot distinguish between genuine collusion, monopsonistic wage-setting, or industry-specific wage norms that happen to cluster at the statutory floor. Indeed, the persistent pre-entry dynamics in the dynamic time fixed effects—with positive and significant coefficients two years before the indicator turns on—confirm that the classification captures a firm type rather than a discrete coordination episode. Future research should exploit sharper quasi-experimental variation in minimum wage policy, such as sector-specific minimum wage thresholds (where they exist), regional variation in enforcement intensity, or discontinuities introduced by discrete minimum wage steps, to establish a causal link between tacit wage coordination and capital structure.

Finally, AdC records the start date of administrative investigation and the formal cartel period established in the decision, which may not coincide precisely with the economic onset of collusion. If the true cartel start predates the recorded start—a common finding in cartel research (Levenstein & Suslow, 2011)—the dynamic time fixed effects may attribute some of the pre-cartel leverage dynamics to the false pre-period, biasing the pre-entry coefficient test toward finding no violation.. The two-year window used in the analysis is also relatively short, implying that cartel activity panels with a larger within-variation component would allow for richer characterization of the full trajectory of leverage adjustment.

8 Conclusions

This dissertation examines the relationship between firm leverage and collusion in Portugal over the period 2010–2022, drawing on three complementary identification strategies: cartel decisions by the *Autoridade da Concorrência* for product and labor market collusion, and the minimum wage spike methodology of Shelkova (2015) for tacit wage coordination.

The results deliver three main findings that together paint a nuanced picture of how financial structure and collusive behavior interact across different market contexts.

For product market collusion, higher leverage is positively associated with cartel formation once firm size, profitability, and market concentration are controlled for. This result survives instrumentation with a Bartik shift-share instrument, confirming that it is not driven by reverse causality. The asymmetric decomposition reveals that the association is stronger among firms accumulating debt than among those reducing it, consistent with the idea that firms taking on new liabilities face sharper incentives to stabilize revenues through collusion. However, once firms are in a cartel, leverage does not fall as predicted by the strategic deleveraging hypothesis of Maksimovic (1988): within the same firm, leverage rises during active cartel spells across all specifications, and the dynamic time fixed effects finds no systematic pre-entry dynamics or post-onset deleveraging. Portuguese product cartels do not appear to use debt reduction as a commitment device.

For labor market collusion identified through AdC enforcement records, the formation results mirror those of product collusion: leverage predicts entry into labor cartels in the same direction and with similar robustness. When the causal direction is reversed and active cartel membership is used to explain leverage, the sign flips: within the same firm, leverage falls significantly during labor cartel spells, consistent with Maksimovic (1988), and the dynamic time fixed effects shows no pre-entry dynamics but a sharp drop at cartel onset. The contrast between the two types of cartel is striking and suggests that the mechanism through which debt interacts with collusion operates differently depending on whether firms are coordinating on product prices or on labor costs.

For tacit wage coordination identified through the minimum wage as a focal point (Shelkova, 2015), firms engaged in minimum wage collusion carry higher leverage than comparable firms, and this association is robust to three alternative operationalizations of the collusion indicator across all fixed-effect structures and estimators. This finding runs counter to both the theoretical prediction underlying H5 and the existing literature, which suggests that wage suppression should increase internal cash flows and facilitate debt reduction. The dynamic time fixed effects reveal that firms classified as minimum wage colluders already carried higher leverage two years before the indicator turns on, with positive and significant pre-trend coefficients. These results should therefore be interpreted as descriptive associations rather than causal estimates: firms that systematically pay workers at the minimum despite their productive characteristics tend to be persistently more leveraged, but the direction of causality cannot be established with the available identification strategy.

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Appendix

Table 21: Endogeneity Test – Leverage and Cartel Formation

	Product Market Cartel		Labor Market Cartel	
	Controls + Partial (Just-identified)	Controls + Partial (Overidentified)	Controls + Partial (Just-identified)	Controls + Partial (Overidentified)
Leverage$_{t-1}$	0.0263^{***}	0.0260^{***}	0.0033^{**}	0.0032^{**}
	(0.0101)	(0.0100)	(0.0015)	(0.0015)
Year FE	✓	✓	✓	✓
Sector FE	✓	✓	✓	✓
Controls	✓	✓	✓	✓
Endogeneity test $\chi^2(1)$	5.143	1.168	5.120	2.721
Endogeneity test p-value	0.023	0.280	0.024	0.099
Hansen J p-value	–	0.110	–	0.251
Observations	3,875,211	3,875,211	3,875,211	3,875,211

Notes: Dependent variable is Cartel$_{it}^{PM}$ in columns (1)–(2) and Cartel$_{it}^{LM}$ in columns (3)–(4). The endogeneity test is the Durbin–Wu–Hausman test implemented via *ivreg2* with the *endog()* option. The *partial()* option removes the contribution of the control variables before estimation, isolating the coefficient on leverage from the exogenous component of the instrument after conditioning on firm characteristics. Just-identified specifications use the employment-share Bartik instrument alone; overidentified specifications add the sales-share Bartik as a second excluded instrument. H_0 (endogeneity test): Leverage is exogenous; rejection in the just-identified case confirms that OLS estimates are inconsistent and IV estimation is required. H_0 (Hansen J): all instruments satisfy the exclusion restriction (are valid); failure to reject confirms instrument validity in the overidentified case. The absence of rejection in the overidentified endogeneity test is consistent with reduced test power from the inclusion of a weaker second instrument; the sign and significance of the leverage coefficient are maintained across all specifications. Standard errors clustered at the firm level. Year and sector (CAE-2 digit) fixed effects included. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.