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Regulation by Public Options: Evidence from Pension Funds*

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Abstract

Many developing countries have private firms administering and investing workers' savings. By relying on private firms, governments protect funds from short-term fiscal needs at the cost of potentially creating market power for these firms. The current political debate in many countries revolves around the role of the high fees charged by private firms in explaining low replacement rates, and the possibility that public options can solve this problem. In this paper, we analyze the equilibrium welfare effects of using public options to regulate market power in the market for individual capitalization pension systems. We develop and estimate a dynamic model of demand and supply in the market for pension funds in Uruguay. We find that the presence of a public option reduces equilibrium fees. Replacing the public option with a private one would increase the fees of the private firms by 16% on average and would more than double the fee of the substitute private option. Additionally, price regulation based on the public option would almost eliminate firms' rents.

JEL Classification: L51, N2, H4, L21

Key words: competition, state-owned firms, pension funds, regulation

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1 Introduction

Many developing countries reformed their traditional pension systems to introduce private firms responsible for administering and investing workers' savings[1]. Private firms have been employed as a commitment device to safeguard employees' long-term savings from the short-term financial needs of the government. Historically, access to these funds or high inflation was utilized to expropriate these savings. Through these reforms, the government tightened its hands at the cost of potentially creating market power for private firms. The current political debate in many countries centers around the responsibility of the high fees charged by private firms in the low replacement rates for employees and the possibility that public options can solve this problem. In this paper, we examine the welfare effects of employing public options as a means to regulate market power in the individual capitalization pension systems market.

The argument in favor of introducing public options is that they can compete more aggressively with private firms by charging lower fees, thus potentially increasing workers' savings. However, while the participation of public options can contribute to solving market failures, existent literature also shows how the equilibrium welfare effects of this policy are *a priori* uncertain Kang (2022). Market segmentation, price-increasing competition, and inefficient provision (Hastings et al. (2017), Chen and Riordan (2008), Duggan and Scott Morton (2006)) can have adverse effects on the welfare of market participants rather than improving it (Jiménez-Hernández and Seira (2021), Atal et al. (2021), Fonseca and Matray (2022)). Therefore, the potential for public options to enhance welfare, as well as the optimal regulation of these markets, represents an empirical question.

In this paper, we study the welfare consequences of the participation of a State-Owned-Firm (SOF) in the Uruguayan market of Pension Fund Administrators (PFAs). In this market, there are 3 private firms and a public option competing to manage workers' savings, with tight regulation on investment portfolios imposed by Law. PFAs charge workers a fee as a share of gross wages to administer their monthly contributions until retirement and are not allowed to charge different fees to different workers. Furthermore, workers enroll through sales force agents and, once enrolled, switching between firms is infrequent. We observe and analyze three different equilibria, motivated by changes in the

[1]Chile (1980), Perú (1993), Colombia (1994), Argentina (1994), Uruguay (1996), Bolivia (1997), México (1997), El Salvador (1998), Costa Rica (2001), República Dominicana (2003), Nicaragua (2004), Ecuador (2004).

SOF shareholders' preferences and the introduction of price caps.

We use detailed administrative data from the institution that administers social security in Uruguay (BPS). We use a panel of a representative sample of workers who made enrollment decisions between the inception of the system in 1996 and 2020 to recover enrollment preferences. We observe monthly data on gross wages, basic employers' characteristics, spells in and out of the formal labor market, PFA enrollment decisions, and mechanism of affiliation, plus basic demographics (sex and date of birth). We complement these data with monthly publicly available market-level data published by the Central Bank about market shares, fees, investment returns, switchers, sales force agents, contributions, and PFAs' financial statements.

We empirically study three specific periods with different regulatory environments and preferences of the SOF PFA shareholders, which we argue gave rise to three different steady-state equilibria. First, we consider a period where the public option sets fees close to those of the private PFAs and without any regulation on fees. Second, we consider a later period where the SOF reduced the administration fee significantly following what we argue is a change in shareholders' preferences but still without fee regulations. Finally, we study a recent period where a cap on fees that affected private PFAs was imposed by the Government.

We develop and estimate a dynamic model of demand and supply in the market for pension funds. We use a demand system in the tradition of (Berry, 1994; Berry et al., 1995). We utilize micro-level data on workers' choices since the inception of the system, and the fact that in this context, the fees are set nationally by firms but the cost of choosing a specific PFA varies across workers, to identify preferences (Hastings et al., 2017). We consider effective administration costs paid by workers, investment returns, PFAs' fixed effects, and idiosyncratic shocks as the key drivers of choices. The administration costs are based on workers' yearly gross wages and observed fees at the time of enrollment. Following the low switching rates observed for already enrolled workers, we assume that switching costs are infinite so only initial choices matter.

On the supply side, we develop a dynamic model of forward-looking single-product firms that compete for new enrollees with no possibility of price discrimination between new and old cohorts of workers. Private firms maximize the present discounted value of economic profits while the public option considers both profits and workers' welfare in its objective function. Given that firms have a high share of old consumers already affiliated with high switching costs and in every period compete for a small cohort of new consumers, in the model firms face a trade-off between investment and harvesting motives (Beggs and Klemperer, 1992; Farrell and Klemperer, 2007). We argue that the fees that we observe are an equilibrium between the two.

We estimate the demand and recover the primitives in each period by grouping workers based on demographics and on whether they had the outside option of remaining not enrolled or not[2]. In general, the estimates show how higher income and older workers are more sensitive to administration fees, results that can be rationalized given the available evidence about how financial literacy is positively correlated with age and income (Lusardi, 2008).

Once we recover preferences, we use them to back out enrollment marginal costs assuming that PFAs compete for new enrollees and set fees to maximize an objective function. Different from the previous literature that studies this market assuming zero marginal costs (Hastings et al., 2017; Luco, 2019; Illanes, 2016), we back out marginal costs and find that they are positive and aligned with the variable payment per enrollee that sales force agents receive from PFAs. However, in the case of the public option, marginal cost can not be separately identified from the conduct parameter that weighs profits and workers' welfare. To disentangle this, we use the range of marginal costs estimated for private firms and assume that they are not significantly different. Using this assumption, we are able to recover the profit and non-profit motives parameters for the public option. We find that the conduct parameter that weights profits in the objective function of the public option effectively decreased between the first and second equilibrium in our sample.

We then proceed to analyze counterfactuals to quantify the value of the regulation by a public option. Using the demand primitives, estimated marginal costs, and the SOF weights between profits and workers' welfare, we investigate welfare changes if the SOF PFA is replaced with a private firm. We find that if there were four private firms in the market, fees of the private firms would increase by 16% on average and the fees of the privatized firm would more than double. Therefore, the presence of the SOF PFA reduces fees and it not only benefits its enrollees directly but also private PFAs' enrollees indirectly through competition.

[2]In Uruguay the retirement system is dual, it includes a pay-as-you-go component in addition to the capitalization one. Enrollment in the latter is mandatory only for workers above an income threshold.

Additionally, we consider a counterfactual to measure how close the with and without regulated fees equilibrium is to an optimal benchmark. In our analysis, we define the optimum as a scenario in which a social planner regulates fees to maximize consumer welfare while ensuring the presence of all 4 active firms in the market. In this case, we assume that the firms preserve their varieties from the workers' perspective, but can charge a fee only enough to recover operational costs. Compared to this benchmark, the public option and the regulation on fees reduce firms' economic profits almost to the minimum that would be possible under the current market structure with 4 firms, while the public option alone without regulation on fees closes that gap between the situation with 4 private firms and the optimum by 50%.

In this paper, we contribute to two strands of the literature. First, we contribute with the empirical literature that analyzes the welfare effects of State-Owned-Firms (Fonseca and Matray (2022),Jiménez-Hernández and Seira (2021), Atal et al. (2021), Handbury and Moshary (2021), Curto et al. (2019),Busso and Galiani (2019), Cunha et al. (2019)). In this case, we contribute by understanding what is the value of the public option in the context of forward-looking single-product firms with market power when the public option has a similar quality as private ones. Our results show that its welfare benefits reach workers enrolled in it as well as those enrolled in private PFAs, due to increased competition and lower fees (Jiménez-Hernández and Seira (2021)).

Second, we contribute to the literature that studies the market regulation of pension fund administrators in individual capitalization retirement systems (Hastings and Tejeda-Ashton (2008), Hastings et al. (2017), Illanes (2016), Luco (2019)). In this case, unlike previous papers, we can observe a public option operating in the market, instead of estimating its effects as a counterfactual (Hastings et al. (2017)). Furthermore, the changes in the regulatory environment and shareholders' preferences allow us to understand how the welfare effects of the public option change when the institutional configuration also changes. The fact that we consider a mature market, with a low proportion of new affiliates in relation to the old ones, in a scenario where switching costs are high, also implies that the effects of the public option are different than what was previously estimated in the literature. A lower fee of the public PFA pushes private PFAs to reduce fees to compete for new enrollees, instead of increasing them to make more profits out of the already enrolled workers.

2 Institutional background

In 1996, Uruguay reformed its retirement system and introduced an individual capitalization component to complement the ongoing pay-as-you-go system. While the latter is administered by the Banco de Previsión Social (BPS)[3] and it is mandatory for every worker[4], the former has two elements. Before retirement, workers' contributions to individual savings accounts are administered and invested by Pension Fund Administrators (PFAs) regulated by the Central Bank (BCU). After retirement, an annuity based on workers' savings is paid by an insurance company until the deceased.

2.1 Workers contributions and PFA enrollment

The workers' contribution rate is fixed at 15% of gross wages[5]. Workers with gross wages above US$ 1,535[6] have to contribute to the individual capitalization sub-system. Conditionally on not being already enrolled the first time they cross this threshold, they have 2 months to choose a PFA when they do so. If they do not choose one, the BPS assigns the worker one by default[7]. Except for this mechanism, workers can only enroll through the sales force agents employed by PFAs. On the other hand, enrollment for workers with gross wages below US$ 1,535 is optional, and if they decide to enroll, they contribute half of their contribution to each subsystem.

Regarding the distribution of savings between sub-systems for individuals with wages above US$ 1,535, for gross wages above US$ 2,303 contributions to the pay-as-you-go sub-system are capped at 15% of US$ 1,535. Therefore, 15% of any dollar over this later threshold and up to a total gross wage of US$ 4,605 goes to individual savings accounts. For gross wages below US$ 2,303, contributions based on gross wages up to US$ 1,535 are divided in half between the sub-systems, and 15% of every dollar earned more than US$ 1,535 and up to US$ 2,303 goes to the pay-as-you-go sub-system.[8]

[3]Equivalent to the Social Security Administration.

[4]Except for small groups of workers that have particular retirement systems.

[5]Employers' contributions go exclusively to the pay-as-you-go pillar.

[6]Converted from Uruguay Pesos to US Dollars using April 2021 reference values. The value of the thresholds is adjusted yearly according to the Average Wage Index.

[7]Until 2014, the assignment was made by lottery in proportion to firms' market share. Since then, they are distributed between the two firms with the lowest administration fees, unless the gap between the two exceeds 20%. If this occurs, default affiliations are made only to the firm with the lowest fee. In practice, this enrollment mechanism accounts for 10-12% of new enrollments each year and since 2014 the SOF PFA has been the only beneficiary.

[8]There is a second alternative to distribute contributions once enrolled but in practice, it is rarely chosen

2.2 PFAs' market structure and regulation

There are currently 4 active PFAs in the market, 3 of them are private firms and the remaining one is a public option: a SOF firm that operates under the rules of private firms but whose shareholders are other state-owned institutions. This option has existed since the inception of the system in 1996 and its shareholders are BROU (SOF bank, 51%), BPS (SSA, 37%), and BSE (SOF insurance, 12%). In the case of private firms, the market started with 5 firms and reached its current equilibrium following the merger of 4 firms into 2 in 2001.

PFAs receive workers' net contributions to individual accounts and invest them in available assets that belong to certain categories authorized by law. For their services, PFAs charge workers a fee that is a percentage of the gross monthly contribution that they make to their accounts in the individual capitalization sub-system. While until 2008, firms could charge both a fixed and a variable fee over the monthly gross contribution, nowadays they can only charge a variable fee. PFAs are not allowed to price discriminate, so they charge every enrollee the same fee. Moreover, in 2018 the Parliament imposed a cap on fees based on the lowest available fee in the market[9], which was fully implemented in 2020 following a transition period of 2 years.

Regarding portfolio rates of return, different regulations in place leave little room for differentiation between PFAs. The fact that close to 60% of total assets are some form of Government debt and that less than 15% of those assets are invested outside the country gives firms limited investment opportunities. Additionally, by law, a PFA is responsible for compensating workers when the rate of return that it obtains is below the minimum between 2% and the average return of the system minus 200 b.p. Furthermore, different from other individual capitalization systems, in Uruguay workers can not decide how to allocate savings between the 2 available investment funds within a PFA. Until 2014 there was a single investment fund available within a PFA, and since then allocation rules between funds are based only on enrollees' age. Taken together, portfolio regulations, workers' insurance, and lack of decision-making regarding risk profiles give PFAs incentives to hold similar investment portfolios and therefore translate into PFAs offering similar rates of returns to their enrollees.

by individuals. It implies that contributions to the pay-as-you-go component are capped at 15% of US$ 1,535, so workers save in their individual accounts 15% of gross wages over that threshold.

[9] The cap stands at 1.5x times the lowest fee in the market.

Finally, workers can switch between PFAs but switching rates have been historically low. To switch PFAs workers need at least 6 months of contributions in the old one, and different from the initial enrollment mechanism, they can not switch through a sales force agent but need to carry out a face-to-face procedure at the PFA office. Though this last requirement was slightly simplified in previous years, switching rates are still relatively low compared with other Latin American countries with capitalization systems.

3 Data and descriptive statistics

3.1 Data

We combine several data sources about workers' characteristics and choices in the pension system, as well as PFAs' financial statements and data about fees, sales force agents, and portfolio rates of return. First, we use a novel database with administrative records collected by BPS for a representative random sample of workers since the inception of the market in 1996 and until 2020. The dataset is a monthly panel of workers' records with information about wages, employer characteristics, demographics (date of birth, sex, area of residence), and PFA enrollment decisions (enrollment mechanism and date, selected PFA, and an ever-switched indicator). Second, we use publicly available information at the market level published by the Central Bank about market shares, fees, investment returns, contributions, switchers, and sales force agents for the period 1996-2022. Finally, we use firms' financial statements for the period 2001-2020, also publicly available on the web of the regulator.

3.2 Descriptive statistics

The markets that are created around the individual capitalization retirement systems have two subsequent phases before the moment when workers start leaving the labor market. At the very beginning, most workers have not chosen a PFA, and therefore there are few workers already enrolled and many to enroll. A model for firms' competition in this period is proposed in Hastings et al. (2017). Then, during the following years, there is a sizeable stock of workers already affiliated and a smaller flow of them entering the market each year. In this paper, we focus on this later phase, which we identify as starting in 2002 when the ratio of new to old enrollees fell below 10%.

In Table 1 we show descriptive statistics at the aggregated market level for the average year of the sample. The average number of already affiliated workers is 1,146,540, while the average size of the cohort of new enrollees entering the formal labor market is 63,317, which represents 5.5% of the stock. Additionally, the average number of switchers per year is low. It is worth noticing that the switching rate (0.31%) is even lower than the ones observed in similar regimes in Latin American countries[10]. Low switching rates in this market have been documented and analyzed in previous work (Luco (2019), Illanes (2016)).

Table 1: AGGREGATED MARKET DESCRIPTIVE

	Average year	% Affiliates
Affiliates	1,146,540	
New affiliates	63,317	5.52%
Net switchers	3,497	0.31%

Notes. Av. for period 2002-2020. Net switchers are expressed in annual terms.

During the period analyzed in the paper, we observe what we consider to be three different equilibria. These equilibria originated in structural changes regarding the preferences of the shareholders of the public option following a change in the Government in 2005 and in the introduction of a cap on market fees[11]. The change in shareholders' preferences materialized in lower fees, and its consequences are easily noticeable in accounting profits. It is worth noticing that, making references to the evolution of accounting profits does not imply that economic profits are not the relevant measure of profits. In fact, it is economic profits the measure that guide the analysis in the following sections. However, given that accounting profits changed significantly during the period, we believe it is worth highlighting them to better identify the three different equilibrium phases we are going to characterize in our supply and demand models.

- **Equilibrium 2002-2005: Relatively high SOF PFA fee, no fee regulation.** This period is characterized by the SOF PFA charging slightly lower fees than private firms and Return-Over-Equity (ROE) similar to the average ROE of private firms.

[10]See https://www.aiosfp.org/ for detailed information on the percentage of switchers over affiliates by country.

[11]These equilibria are depicted in grey in Figure 1.

- **Transition 2006-2013: SOF PFA reduces fees, no fee regulation.** In 2006, the SOF began to reduce its fee, a change of behavior that we argue comes from a change in the preferences of its shareholders, as documented in public shareholders meetings' minutes. The change had the purpose of favoring workers and allowed the PFA managers to reduce the fee as long as the ROE remained above some minimum threshold level imposed by the main shareholder[12].

- **Equilibrium 2014-2017: Low SOF PFA fee, no fee regulation.** The SOF continued with the same mandate from the shareholders, but the ROE became a constraint to continuing the fee reduction policy. If we compare this equilibrium with the first one, the fee of the SOF PFA fell by half while private PFAs reduced it only slightly. As a result, during this period the ROE of private PFAs increased and that of the public option fell from 40% to 12%. This is preliminary and descriptive evidence of a change in the behavior of the SOF PFA. However, in our model, this behavioral change of the SOF will not be imposed but recovered from the data.

- **Transition 2018-2019: Low SOF PFA fee, progressive implementation of a cap on fees.** Between 2018 and 2019, the regulator implemented a transition phase that allowed firms to converge from pre-regulation observed fees towards the maximum ones allowed by the new law: firms were allowed to charge workers a fee up to 50% above the lowest fee in the market. The discussion about the potential introduction of the cap began in December 2017, so it doesn't affect our analysis of the previous period. The reduction in PFA fees observed during this period is motivated by this regulation. Since 2019, the fees of private PFAs have been equal to the maximum established by law.

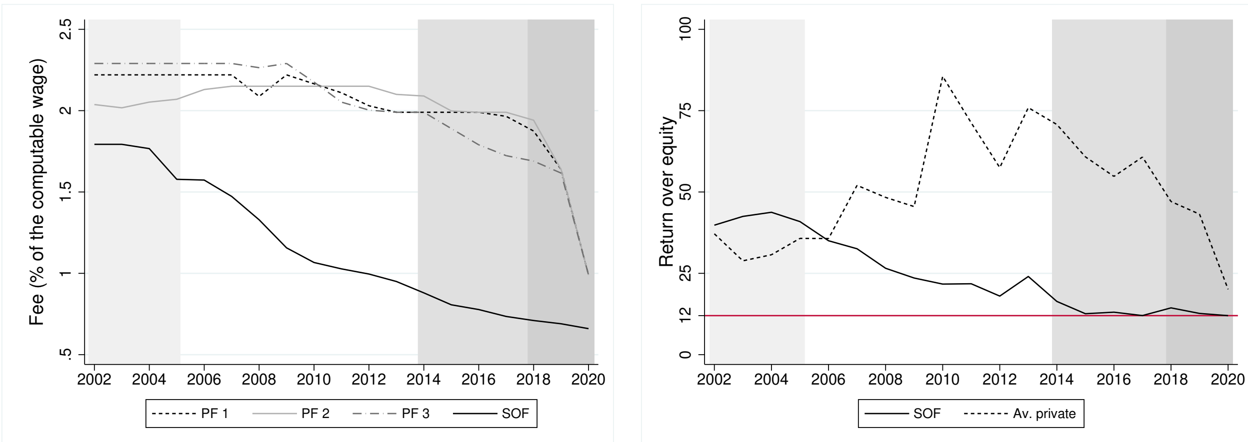
- **Equilibrium 2020: Low SOF PFA fee, fully implemented cap on fees.** The fully implemented policy of caps on fees generated a new equilibrium in which private PFAs' fees fell by half with respect to the average 2014-2017. Additionally, private PFAs ROE decreased from approximately 60% to 20%.

Within each equilibrium period, private PFAs' fees are relatively stable and have no sizeable difference between them. Additionally, in Figure A.1 it can be observed that the

[12]Shareholders' meeting for the fiscal year 2017: "(...) *distributable profits stood at a ROE of 14.4%, exceeding the minimum requirement of 12% established by the majority shareholder.*" (...) "*it is requested to continue with the fee reduction policy (...)*".

Figure 1: EVOLUTION OF FEES AND ROE OVER TIME

(a) Fees (b) ROE



Note. Fees are reported as a percentage of the gross wage fraction relevant for contributions to each subsystem. Return-Over-Equity is calculated as the ratio between distributed profit in t and equity in $t-1$. The red line represents the minimum acceptable ROE imposed by the main shareholder on the SOF PFA. Shaded areas indicate the 3 equilibrium periods: 2002-2004, 2014-2017, and 2020.

evolution of the share of old workers does not show important changes over time. The SOF PFA is the leading firm with almost 40% of the market. There is also heterogeneity in shares across the wage distribution, with the SOF PFA obtaining a greater market share the higher the wage level of the worker[13]. In fact, for low-wage workers, the public option is currently not the leading firm in the market, something that looks consistent with previous evidence on individual capitalization pension systems and on the correlation between income and financial literacy (e.g. Hastings and Ashton (2008); Lusardi and Mitchell (2011); Hastings et al. (2017)).

To complete the description of the firm's characteristics, in Figure A.2 we show the evolution of the real annual rate of return by firm and the difference with the market average. There is no evidence of persistent out-performance of private PFAs over the SOF PFA. On the contrary, the median return of the public option during the entire period is 4.1%, 10b.p. above the median of private PFAs. Additionally, the median rate of return difference between the firms ranked in second and third place is 30 b.p., which implies that similar to what happens with similar systems in other Latin American countries, the rate of return is not a first-order dimension of the competition and differentiation between firms (e.g. Luco (2019) for Chile and Hastings et al. (2017) for Mexico). About the sales

[13]See shares by income bracket in Table A.1 in the Appendix.

force, in Table A.2 of the appendix we show the average share by firm in each period. It can be seen that the SOF PFA is the leader also in terms of sales force agents, with an average between 35% and 36%, and their share is stable between the three periods.

In Table 2 we show descriptives of selected demographics for the microdata sample. We have 125.453 individuals, of which 48% are women. The median age of the individuals entering the formal labor market is 23.5 years old, and the median age of enrollment is 24.9. In Figure A.4 it can be seen that conditional on enrollment and for non-forced enrollees, 75% do so in the first two years after entering the market. In the main demand specification, we work with all individuals, regardless of the time that elapses between entering the labor market and their affiliation[14]. The median enrollment wage in the sample is US$ 834 and 15% of the sample had gross wages above the first contribution threshold at the moment of entering the market. For its part, for those individuals whose salary ever exceeded the threshold that requires them to enroll, 24% were enrolled by default. Finally, for individuals that have the option not to enroll (gross wages below the first income threshold), 74% choose to save in an individual account in the capitalization sub-system.

Table 2: WORKERS' SAMPLE SUMMARY STATISTICS

Individuals	125,453
Gender (female)	0.48
Age when entering the market (median)	23.2
Age when enrolling (median)	24.9
Gross wage (median, US$)	834
Share with enrollment gross wage above threshold (US$ 1,535)	0.15
Outside option (conditional on gross wage below US$ 1,535)	0.26

Notes. The Table reports descriptive statistics for selected demographics for the available sample. Average 1996-2020. UYU expressed in US$ 2017.

4 Model

The model characterizes the competition of forward-looking firms in the individual capitalization pension market. As described before, the market is composed of single-product

[14]As a robustness check, we conduct an alternative demand estimation only with individuals who enter the market and make their first decision in the first two years after entry, with no relevant differences in the pattern of the elasticities.

firms, with no possibility of price discrimination between new and old cohorts of workers. Private firms have similar and stable prices and market shares over time and workers present a behavior consistent with high switching costs. In markets with these characteristics, firms face a trade-off between investment and harvesting motives. Following Beggs and Klemperer (1992) and Farrell and Klemperer (2007), we work using the fact that in these markets exists an equilibrium with constant prices over time. This price is higher than the equilibrium price that would emerge in a market without switching costs.

The main focus of the model is to capture the investment harvesting trade-off and the strategic interaction of private firms with the public option. On the supply side, we model single-product forward-looking private firms competing in prices. The SOF firm has an objective function that includes both profits and workers' welfare. On the demand side, the model represents workers' choices among firms for cohorts entering the market, and we assume that they face infinite switching costs.

In our model, we make the following assumptions. First, workers choose a PFA considering variables at time t and after that remain enrolled in the firm until retirement. This is consistent with the reduced levels of switching observed in the market. Second, following the previous literature on dynamic competition, we assume that firms play a dynamic game. The previous assumptions imply that firms are forward-looking but consumers are myopic[15] and have infinite switching costs. Third, given that in each equilibrium we do not observe meaningful variation in fees, we assume that firms play a closed-loop game, committing to a single fee during the entire period. Fourth, we focus on price competition and rely on the fact that, given the regulation, firms cannot substantially differentiate in rates of return. Then, we assume that firms' non-price characteristics are set exogenously. Finally, we allow the SOF PFA to have not-for-profit motives as a device to incorporate the mandate to reduce the administration fee that its SOF stakeholders impose on it.

We model a simultaneous decision game for all firms, with the following timing regarding the occurrence of events. First, firms simultaneously set fees at t for all periods (from t to T), given an initial share of old workers, workers' preferences, and the state of the institutional environment (SOF PFA shareholders' preferences and fee regulation). Second, individual idiosyncratic shocks at time t realize, new workers entering the labor market choose a PFA to remain enrolled there until retirement R years later, and a fraction of old workers retire.

[15] Their discount factor is equal to zero.

4.1 Demand

We model the demand of new workers d_{ijt} with a conditional logit specification where individual i chooses between PFAs to manage her savings until retirement. The indirect utility that worker i receives from firm j is:

$$u_{ijt} = \theta_{it} \times C_{ijt}(y_{it}, f_{jt}) + \delta_{jt} + \epsilon_{ijt} \tag{1}$$

The term C_{ijt} represents the cost of administration that individual i has to pay to firm j in year t, which depends on the worker's short-term expected stream of gross wages y_{it} and the fee of firm j at t. The parameter θ_i represents the cost sensitivity, which, as explained in greater detail in equation 7, we will allow varying linearly with individual's i gross wages at t. The term δ_{jt} considers non-cost components of firm j, some of them observed (rates of return r_{jt}) and others unobserved by the econometrician (brand value ξ_{jt}). Finally, we assume that ϵ_{ijt} is drawn i.i.d. from a Type 1 Extreme Value distribution.

4.2 Supply

In this section, we develop the supply side model that describes the competition in fees of forward-looking single-product firms. In every period, these firms face two types of workers. On one side, there are new cohorts of workers entering the formal labor market. These individuals make a one and for all decision based on product and PFAs' characteristics, as described in the previous section. On the other side, there are already enrolled workers (old) with infinite switching costs, that remain in the same PFA until retirement.

Firms face the investment-harvesting trade-off. On one hand, firms have incentives to harvest by charging a high fee to their existing base of workers to make more profits at t. On the other hand, by charging lower fees, PFAs can obtain more market share at t of new cohorts of workers that are going to remain enrolled between t and retirement at $t + R$, and therefore, to make more profits out of them in the future.[16]

4.2.1 Current period profits function

We first describe the per-period profits function and then, the net present value function for the firms. The per-period profits function of firm j in period t is:

[16]For a more detailed version of the supply model and the computation of the equilibrium fees, see Appendix C.

$$
\begin{aligned}
\pi_{jt} &= \underbrace{f_j \times M_t \times (\underbrace{s_j^n(\mathbf{f}) \times \alpha \times (1-\rho_t^n)}_{\text{New workers}} + \underbrace{s_j^o \times (1-\alpha) \times (1-\rho_t^o)}_{\text{Old workers}})}_{\text{Revenue}} \\
&- \underbrace{MC_j \times s_j^n(\mathbf{f}) \times \alpha \times M_t}_{\text{Enrollment Cost of New Workers}} - \underbrace{F_j}_{\text{Fixed cost}}
\end{aligned} \tag{2}
$$

The terms f_j and M_t represent the fee of firm j and the total mass of gross wages affected by the individual capitalization sub-system in period t, respectively. The shares $s_j^n(\mathbf{f})$ and s_j^o of new and old workers are compound terms that depend both on the workers' enrollment decisions in current and previous periods, as well as on weights associated with the importance of individual gross wages in the aggregated wage mass M_t. This reflects the fact that richer individuals have a higher positive impact on PFAs' revenues than poorer ones through the one-to-one connection between revenues and gross wages. Also, notice how the share of new workers depends on the vector of fees while the share of old ones does not, following the infinite switching cost assumption. Additionally, α is the share of M_t associated with new workers entering the labor market, ρ is the share of retirees, and F_j firms' fixed costs.

Furthermore, instead of assuming marginal costs equal to 0 as in previous literature (Hastings et al., 2017; Luco, 2019), following anecdotal evidence we gathered from industry participants we are going to differentiate between the marginal cost of enrolling a new worker and the marginal cost of managing an additional account. While we still work under the assumption that the latter is irrelevant to the pricing problem, we are going to estimate the former in the model using the Nash-Bertrand competition assumption. Therefore, in our model, MC_j represents the variable payment that PFAs make to their sales force agents for affiliating new workers. More specifically, is the cost per dollar of wage. This variable payment is usually tied to the "quality" of the new enrollee in terms of wage level, expected density of future contributions, etc. To better reflect this feature, the effective marginal cost of enrolling an additional worker depends on the income characteristics of new enrollees.

4.2.2 Net Present Value profits function

Now we discuss the net present value of the profits function, assuming that firms are committed to a single price in all periods from t to T. The net present value of firm j in period t expressed as a summation of cohorts is:

$$\begin{aligned} V_{jt} &= f_j \times M_t \Big[W^o(\alpha,\beta,\rho_t^o) \times s_j^o + W^n(\alpha,\beta,\rho_t^n) \times s_j^n(\mathbf{f})\Big] \\ &- \frac{(1-\beta^T) \times MC_j \times \alpha \times M_t \times s_j^n(\mathbf{f})}{(1-\beta)} - \frac{(1-\beta^T) \times F_j}{(1-\beta)} \end{aligned} \tag{3}$$

where the weights between old and new workers are given by the following expressions:

$$\begin{aligned} W^o &= (1-\alpha) \times \Big[(1-\rho_1^o) + ... + (1-\beta^{t-1}) \times \Pi_{i=1}^{t}(1-\rho_i^o) + ... + (1-\beta^{T-1}) \times \Pi_{i=1}^{T}(1-\rho_i^o)\Big] \\ W^n &= \alpha \times \Bigg(\underbrace{\sum_{t=1}^{t+R} \beta^{t-1}}_{\text{Cohort 1}} + \underbrace{\sum_{t=2}^{t+R} \beta^{t-1}}_{\text{Cohort 2}} + ... + \underbrace{\sum_{t=T}^{T} \beta^{t-1}}_{\text{Cohort T}} \Bigg) \end{aligned}$$

with R being the expected years of work until retirement.

From Equation 3, the problem is similar to the static one in t. However, the value function has a different weight for new and old workers than the static problem. These weights depend on α, ρ_t^o, ρ_t^n, R and β. We assume $\beta = 0.99$ and set α according to the observed relative size of new cohorts of workers entering the labor market. The parameter ρ_t^o is set according to the observed age composition at each period t and assumes that the age of retirement is 62.5 (average retirement age in Uruguay). Finally, we assume that new workers retire 41 years after enrollment ($R = 40$) and that the wage mass M_t remains constant.

We assume the existence of a pure-strategy Nash-Bertrand equilibrium in fees, and that the fees that support it are strictly positive (and lower than 100% of the contribution). In this context a Nash-Bertrand equilibrium in this game is a vector of fees (f_j), such that

$$(f_j) \in arg\,max_{\{f_{jt}\}_0^T} V_{jt}(\mathbf{f}|W^o, W^n, MC_j) \tag{4}$$

for each PFA $j \in J$.

4.2.3 SOF PFA objective function

We assume that the objective function that the SOF PFA maximizes takes into account two goals: the net present value of profits (as in the case of private firms) and workers' welfare. We measure workers' surplus according to the standard formula derived from the expected utility of a model with T1EV shocks (see Equation (7)):

$$CS_i(\mathbf{f}) = \frac{1}{\theta_i} \times log\left[1 + \sum_j exp(\theta_i \times C_{ij}(y_i, f_j) + \delta_j + \epsilon_{ij})\right] \tag{5}$$

And consider the average $\overline{CS}_{sof}$ in the public option objective function.

The weight of each objective is given by the conduct parameter $\lambda \in [0, 1]$, where $\lambda = 1$ implies that the SOF behaves as a private firm (full for-profit) and $\lambda = 0$ implies that it only considers workers' welfare (full not-for-profit). Therefore, the objective function of the SOF is:

$$\begin{aligned} \mathcal{W}(\mathbf{f})_{soft} &= \lambda \times \underbrace{V(\mathbf{f})_{sof}}_{\text{NPV}} \\ &+ (1-\lambda) \times \underbrace{\left(\overline{CS}_{sof}(\mathbf{f}) \times s^n_{sof}(\mathbf{f}) \times M \times W^n + \overline{CS}_{sof}(\mathbf{f}) \times s^o_{sof} \times M \times W^o\right)}_{\text{Workers' welfare of SOF enrollees}} \end{aligned} \tag{6}$$

5 Estimation and results

5.1 Demand estimation and identification

We estimate workers' demand following Hastings et al. (2017). In order to allow for flexible preference heterogeneity, we estimate conditional logit models separately for 16 demographic brackets (or cells c) using workers' microdata. First, we divide the population between individuals with and without outside option[17]. For individuals without an outside option, we divide the population into 4 cells that consider 2 age groups (below and above the median) and gender. We don't consider income brackets for this group because they are already relatively high-income workers. Finally, for individuals with an outside option we use 12 brackets according to 2 age groups, gender, and gross wage tertiles.

[17] See section 2

$$u_{ijt}^c = (\alpha^c + \gamma^c \times w_{it}) \times C_{ijt}(y_{it}, f_{jt}) + \eta_j^c \times \zeta_t^c + \epsilon_{ijt}^c \tag{7}$$

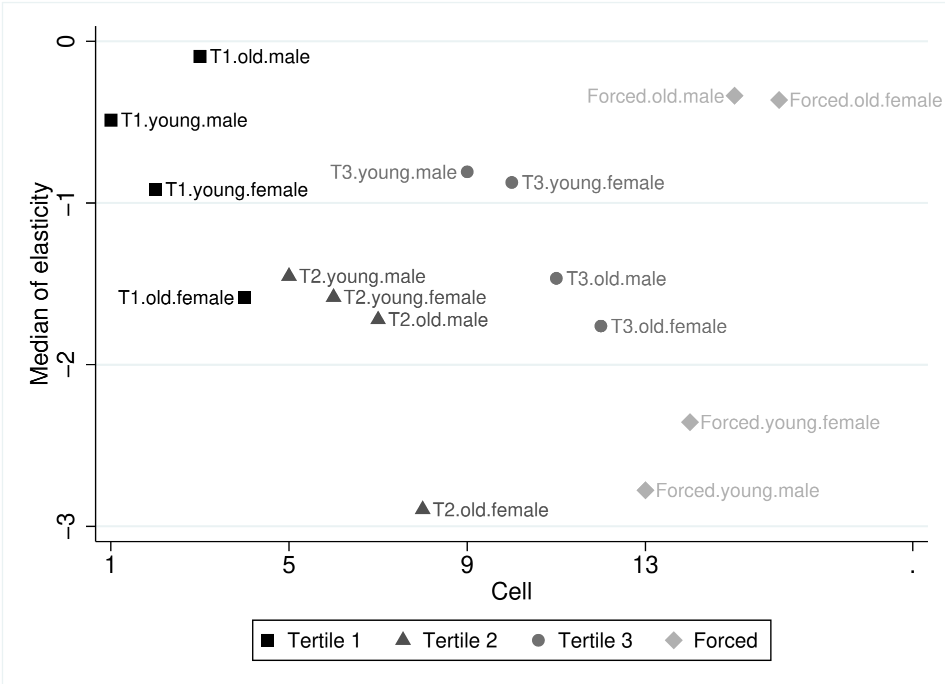
In the demand specification, $\alpha^c + \gamma^c \times w_{it}$ reflects the sensitivity to the administration cost. This sensitivity is allowed to vary linearly with individual i's current monthly gross wage w_{it}. We include interacted fixed effects by firm η_j^c and year ζ_t^c, in order to capture variation in workers' preferences over time in PFAs' non-cost characteristics. To calculate the effective administration cost C_{ijt}, we use the present discounted value of the stream of gross wages during the first 12 months m after individual i enters the formal labor market $y_{it} = NPV(\{w_{im}\}_m^{m+12})$. This is consistent with assuming that workers consider their total administration costs at t.

The identification of the price sensitivity comes from the fact that fees are plausibly exogenous Hastings et al. (2017). In this setting, firms set national fees, but costs are worker-specific and vary with gross wages and spells in the formal labor market. This individual-level variation of the effective administration cost of every PFA, even among individuals with similar demographic characteristics, gives us arguably exogenous variation to estimate the relevant fee sensitivity parameters.

5.2 Demand results

Figure 3 shows the median demand elasticity to the administration cost C_{ij} for individuals in each estimation bracket. We calculate these elasticities using the estimated parameters and observed fees and characteristics. Demand estimates are presented in Tables A.3 and A.4 of the Appendix. Overall, our estimates imply low elasticities, with older and female workers being relatively more elastic. Low-income workers are also more inelastic. This pattern is in line with the previous literature on individual capitalization pension systems in other Latin American countries (e.g. Hastings and Ashton (2008); Hastings et al. (2017)), and on financial literacy, Lusardi (2008). In Figure A.5 we show the elasticities by period, with the same general pattern.

Figure 3: MEDIAN ELASTICITY BY BRACKET



Note. Within bracket median PFA elasticity (all period). Elasticities are calculated at the observed fee levels and individual characteristics. Computed using estimates from equation (7) to generate the logit choice probability for each individual for each firm.

5.3 Supply: backing out of marginal cost

We back out enrollment marginal cost MC_j for two different periods, characterized by the different behavior of the SOF (2002-2005 and 2014-2017) using the demand primitives of each period, the observed shares of already enrolled workers s_j^o and fees f_j, solving the system of first-order conditions (FOC) given by equation (8) for PF and (10) for SOF:

$$M \times [s_j^o \times W^o + s_j^n(f) \times W^n + \frac{\partial s_j^n}{\partial f_j} \times f_j \times W^n] - \frac{\frac{\partial s_j^n}{\partial f_j} \times M \times \alpha \times MC_j}{1 - \beta} = 0 \tag{8}$$

$$\lambda \quad \times M \times [s^o_{sof} \times W^o + s^n_{sof}(f) \times W^n + \frac{\partial s^n_{sof}}{\partial f_{sof}} \times f_{sof} \times W^n - \frac{\frac{\partial s^n_j}{\partial f_j}\alpha MC_j}{1-\beta}] \tag{9}$$

$$+ \quad (1-\lambda) \times M \times \left(\frac{\partial \overline{CS}_{sof}(\mathbf{f})}{\partial f_{sof}} \times s^n_{sof}(\mathbf{f}) \times W^n \right.$$

$$\left. + \quad \overline{CS}_{sof}(\mathbf{f}) \times \frac{\partial s^n_{sof}(\mathbf{f})}{\partial f_{sof}} \times W^n + \frac{\partial \overline{CS}_{sof}(\mathbf{f})}{\partial f_{sof}} \times s^o_{sof}(\mathbf{f}) \times W^o \right) = 0$$

In Table 3 we present marginal costs (expressed as the cost of enrolling an individual with the average gross monthly wage of new workers $MC_j \times \overline{w}^n$) for both periods. For private firms, we estimate enrollment marginal costs between US$ 30 and US$ 50 in 2002-2005, and increasing to US$ 49 - 76 in 2014-2017. As discussed in the model, the marginal cost in this market is mainly associated with paying sales force agents for new enrollees. Although we do not directly observe the variable wage component that sales force workers receive for each affiliation, using data on the sales force average productivity[18] and aggregated wages, and on their minimum wage established in Collective Agreements, we can approximate the observed variable wage they earn. For 2017 this calculation implied a mean variable wage component of US$ 63 dollars per enrollee[19], a figure close to our enrollment marginal cost estimates.

Table 3: Back out marginal cost and SOF profit motive

Period	Marginal cost				Profit motive SOF	
	PF 1	PF 2	PF 3	SOF*	Mean	Min
	(1)	(2)	(3)	(4)	(5)	(6)
2002-2005	45	50	30	-191	0.86	0.86
2014-2017	76	75	49	-403	0.72	0.73

Notes. Back out of marginal cost based on FOC of equation (8). Back out of SOF profit motive based on FOC of equation (10), imposing mean marginal cost of PF in column (5) and the minimum in column (6). For each period, the primitives of demand of that period are used. Marginal cost in US$ 2017. *Assumes full for-profits ($\lambda = 1$).

For the SOF we cannot separately identify the marginal cost and for-profits motives. We then analyze two scenarios. In the first scenario, presented in column 4 of Table 3, we

[18] Calculated as the ratio between monthly new enrollees per firm over monthly total sales force agents.
[19] In the Appendix B we show in detail how we arrived at these values.

recover the marginal cost of the SOF, imposing $\lambda = 1$. Under this full-for-profits behavior of the SOF PFA, we obtain negative marginal costs. Therefore, we cannot rationalize its behavior as a pure profit-maximizing firm, not even during the first equilibrium under consideration before the change in shareholders' preferences. In the second scenario, we impose the average marginal cost of private PFAs on the SOF PFA to recover for-profit motives. Results are presented in columns 5 and 6 of Table 3. Under this assumption, we observe a decrease in the weight of for-profit motive between the first and second period of 0.86 to 0.72[20], consistent with the descriptive evidence and the information we gathered about the SOF PFA behavioral adjustment following a change in shareholders' preferences in 2006.

6 Counterfactuals

We use our estimates about preferences and marginal costs to understand the value of the public option and the effect of the regulation on fees in the market. In particular, we analyze two counterfactuals. In the first one, we substitute the public option with a private firm that resembles the other private options. Here we try to capture the value of the public option by comparing the observed market equilibrium with an alternative configuration with private PFAs only. In the second group, we analyze how far were observed equilibria from an optimal benchmark where fees were set such that PFAs have zero economic profits. We separate the analysis between 2014-2017 and 2020 to account for the effects of the introduction of the fee regulation in the later period. We calculate counterfactual equilibrium fees using the demand primitives of the period of interest, but always with the marginal cost parameters of 2014-2017. In Table A.6 we show the implementation details.

To analyze the impact of these policies, we use measures of firms' profits and workers' welfare, the latter measured by the standard consumer surplus. As discussed in Hastings et al. (2017), given the role that advertising and sales force agents play in enrollment decisions, as well as the lack of knowledge about whether it is an informative or persuasive role, it is debatable whether it is best to perform welfare calculations based on the demand estimates. Therefore, an alternative is to directly analyze changes in PFAs profits, which

[20]The decrease is from 0.86 to 0.73 if we use the minimum marginal cost. Given that the difference generated between both criteria is small, hereafter we work with the average marginal cost for the SOF PFA.

imply a change of equal amount and opposite sign in workers' net savings. From this perspective, the social outcomes that we analyze in the counterfactuals do not imply changes in total welfare, but in its distribution between workers and firms[21]. This is particularly relevant considering the ongoing discussion in Latin America about how should these systems be redesigned to improve workers' outcomes and the trade-off between reducing firms' market power and tying Governments' hands in the use of funds.

Period 2014-2017 During this period there is no cap on fees and the SOF PFA is playing with lower for-profit motives[22]. In the baseline, reported in row 1 of Table 4, we observe how private firms charged workers fees that were more than twice as high as that of the public option.

We present two counterfactuals. The first one addresses a question similar to Hastings et al. (2017): what are the effects on private firms' fees of introducing a public option? Their main result is that this policy generates an increase in private fees as the optimal response. In our setting, however, we start from a baseline situation in which the public option is already participating, and with a mature market where workers' switching costs are key for understanding firms' strategic interactions. In this dynamic competitive setting with high switching costs, the share of already enrolled workers at t is relevant for the equilibrium (the greater the share, the more the harvesting motive weighs). Therefore, we conduct a counterfactual consisting of assuming that the SOF PFA is not in the market, and instead, there is a fourth private firm (row 2 of Table 4), whose characteristics are an average of the 3 private ones observed[23]. We include an additional PF because it is reasonable to think that the non-existence of SOF would lead to the existence of a 4th PF, and because we want to abstract from the effect generated by going from 4 to 3 firms that is not explained by its public option nature.

In the counterfactual with 4 private PFAs, the optimal response of private firms already in the market is to increase fees on average 16%, reducing workers' welfare by 5%. However, given the fact that the fee of the 4th private firm that we introduce raises significantly,

[21] Except for the effect of fees on workers' diversion to the outside option. Our estimates do not indicate that this would have a strong effect on the market though.

[22] See Section 5.3

[23] In Table A.5 of the Appendix we show a second counterfactual, assuming that the SOF PFA plays as a profit-maximizing firm. Interestingly, in this equilibrium, the SOF charges a fee of 5.8% because it can exploit the fact that workers have a very inelastic residual demand for it. Nevertheless, this counterfactual is not the preferred one when it comes to understanding what the market would be like without a public option because we can not disentangle the different elements that compose the non-price characteristics of the public option.

total average fees increase by 30%. We understand that this counterfactual uncovers the value of the public option in a situation where we hold fixed the number of firms in the market. Notice that the positive effect of the public option benefits all workers through the competition for new enrollees, but fundamentally those enrolled in it. With this counterfactual, we show how the presence of the public option helps discipline the market, different from what was estimated in Hastings et al. (2017)

The second set of counterfactuals, reported in row 3 of Table 4 is carried out to understand how close the baseline equilibrium is with respect to an optimal benchmark where there is a social planner whose objective function is to maximize workers' welfare while maintaining the existing varieties in the market by allowing firms to cover their operating costs. Compared to the baseline, the benchmark implies an average reduction of 61% in private firms' fees and an increase in workers' welfare and total welfare of 9% and 1%, respectively. This counterfactual ignores potential economies of scale that would result from the existence of a centralized institution managing savings accounts and the social gains that would result from reducing sales force expenses. This exercise does not consider either how private PFAs incentives would change and affect non-price characteristics, for example, portfolio returns through changes in PFAs' investment policies, or the deployment of sales force agents. We do believe that it is still a useful benchmark to compare the effects of different policies implemented in the market.

Table 4: COUNTERFACTUAL ANALYSIS. PERIOD 2014-2017.

	Fee (%)				Av. Δ	Outside	Profit		CW	ΔCW	TW	ΔTW
	PF 1	PF 2	PF 3	SOF	PF fee	option	PF	SOF		(%)		(%)
1- Baseline	1.98	2.02	1.85	0.80	-	21	41	8	551	-	600	-
2- 4 PFA	2.24	2.31	2.21	1.90*	16	25	106	-	523	-5	629	5
3- Optimal	0.74	0.67	0.40	0.55	-61	15	0	0	600	9	600	1

Notes. Fee as a % of the gross wage component relevant for contributions to individual savings accounts. (2) distribute SOF PFA enrolled workers equally. CS consumer-worker welfare. TW total welfare. CS, Profit, and TW are per year and expressed in US$ millions. (*) Fee of the 4th private PFA.

Caps as Fee Regulation In 2020, the cap on fees is implemented and the SOF PFA plays with low for-profit motives. In row 1 of Table 5 we show the baseline, with the three private firms charging the maximum fee allowed by the regulator (up to 1.5x times the minimum market fee).

In this case, we present two counterfactuals to understand the consequences of the cap on fees. In row 2 of Table 5 we show a counterfactual with the optimal benchmark described previously, but using the demand primitives of 2018-2019[24]. In this case, fees of private PFAs are 39% lower and workers' welfare 5% higher than in the baseline scenario. Additionally, in row 3 of Table 5 we show the fees that would have existed without caps, to understand the effects of the regulation. In this case, fees of private PFAs in the scenario without caps would have been 92% higher while workers' welfare would have been 9% lower. Based on these counterfactuals, the regulation on fees (Baseline 2020) favored workers by reducing fees and increasing their net savings. Without it, private firms would have charged fees almost twice as high for their services.

Table 5: COUNTERFACTUAL ANALYSIS. PERIOD 2020.

	Fee (%)				Av. Δ	Outside	Profit		CW	ΔCW	TW	ΔTW
	PF 1	PF 2	PF 3	SOF	PF fee	option	PF	SOF		(%)		(%)
1- Baseline 20	0.99	0.99	0.99	0.66	-	13	8	2	436	-	446	-
2- Optimal	0.75	0.66	0.40	0.55	-39	12	0	0	456	5	456	2
3- No caps	2.01	1.91	1.78	0.66	92	17	40	2	399	-9	441	-1

Notes. Demand 2018-2019, enrollment marginal cost 2014-2017. CS consumer-worker welfare. TW total welfare. CS, Profit, and TW are per year and expressed in US$ millions.

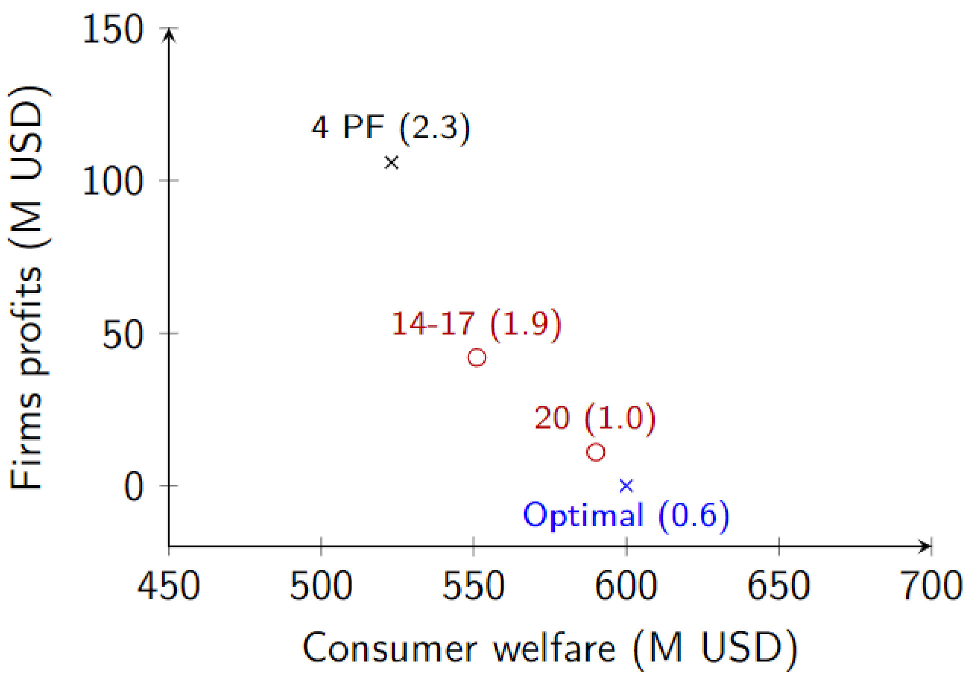
Summary Finally, in order to compare the observed outcomes in 2014-2017 and 2020[25] with the two main counterfactuals, we summarize the results in Figure 4. To abstract from changes in preferences between periods, we present the results using only the demand estimates for 2014-2017. The main conclusion is that, either by looking at workers' welfare or firms' profits, the equilibrium in 2020 where the SOF PFA plays with low for-profit motives and there is a binding cap on fees generated market outcomes that were near the optimal benchmark. Additionally, if we think in terms of how close each institutional arrangement was to this benchmark, the inclusion of a SOF PFA led to profits that were almost half the distance between the market equilibrium with 4 private PFAs and the benchmark. Furthermore, the public option together with the regulation on fees shrank

[24]The sample ends in April of 2020, so we assume that time-firm fixed effects remain constant in 2020. Additionally, as a result, the consumer welfare levels included in this table are not comparable with the figures included in Table 4.

[25]In both cases, the SOF PFA has low for-profit motives, but only in the latter the cap on fees is known by market participants and binding.

the profit gap between the two theoretical equilibria by more than 90%.

Figure 4: SUMMARY MAIN COUNTERFACTUALS



Note. Demand primitives, marginal costs, and for-profits conduct parameter of 2014-2017. Profits and consumers-workers welfare per year expressed in US$ millions. In parentheses, 3 private PFAs' average fees in each equilibrium. 1) 4 PF: replaces SOF with private PFA with average characteristics of the 3 observed. 2) 14-17: baseline for that period, no cap on fees and 3) 20: baseline for that period, cap on fees, 4) Optimal: social planner maximizes workers' welfare s.t. covering the 4 observed firms' operational costs.

7 Final comments

This paper examines the competitive and welfare consequences of the participation of a SOF PFA in a mature individual capitalization pension system in which workers face high inertia and low fee sensitivity. Relying on a data set with rich details on workers' demographics and enrollment decisions, we estimate workers' demand and use those estimates to recover both the marginal cost of enrollment and the weight of for-profit motives in the SOF PFA objective function. Marginal cost estimates are broadly in line with variable payments that sales force agents receive for enrolling workers, while the model pins down a decrease in for-profit motives beginning in 2005 that we document in the SOF PFA shareholders minutes.

We conduct two main sets of counterfactuals. In the first one, we estimate that administration fees of private PFAs are 16% lower (and total fees 30% lower) due to the presence of a public option with not-for-profit motives. Therefore, we show that the presence of a public option helps discipline the market by forcing private firms to compete more aggressively for new enrollees. In the second one, we calculate the optimal regulation for workers, by setting fees in a way that private firms and the SOF have zero economic profits. In this case, optimal average fees of private PFAs are 61% lower than the baseline in 2014-2017 (no fee caps) and 39% than in 2020 (fee caps). Additionally, we show that the baseline situation of 2014-2017 generates profits that are 50% closer to the optimum if we compare with the market solution without a public option, while the situation with the SOF PFA plus the cap on fees brings the market 90% closer to the optimal.

Our findings suggest that creating a public option with high not-for-profit motives helps to discipline the market power of private firms. However, the competitive effect is not strong enough to bring the equilibrium to the optimal point, leaving room for additional regulation. For its part, the combination of a public option with fee regulation seems to be a good alternative to additionally reduce the market power of private firms in a market with little room for differentiation in rates of return, while preserving the ability of this system to discipline Governments by making it more costly for them to use the savings. Nevertheless, our analysis leaves open the question of what efficiency gains can be generated by centralizing the management of individual accounts and allowing private firms to compete only for the investment of funds.

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A Appendix

Table A.1: Market shares of new enrollees by wage bracket and period

	Firm			
	PF 1	PF 2	SOF	PF 3
Period 2002-2005				
Wage tertile 1	0.21	0.27	0.19	0.32
Wage tertile 2	0.18	0.25	0.31	0.27
Wage tertile 3	0.17	0.17	0.48	0.18
% above threshold	0.12	0.10	0.64	0.14
Period 2014-2017				
Wage tertile 1	0.25	0.28	0.21	0.26
Wage tertile 2	0.21	0.23	0.26	0.30
Wage tertile 3	0.16	0.22	0.36	0.27
% above threshold	0.10	0.15	0.57	0.18
Period 18-19				
Wage tertile 1	0.24	0.24	0.23	0.28
Wage tertile 2	0.20	0.22	0.24	0.34
Wage tertile 3	0.21	0.17	0.33	0.29
% above threshold	0.08	0.02	0.85	0.05

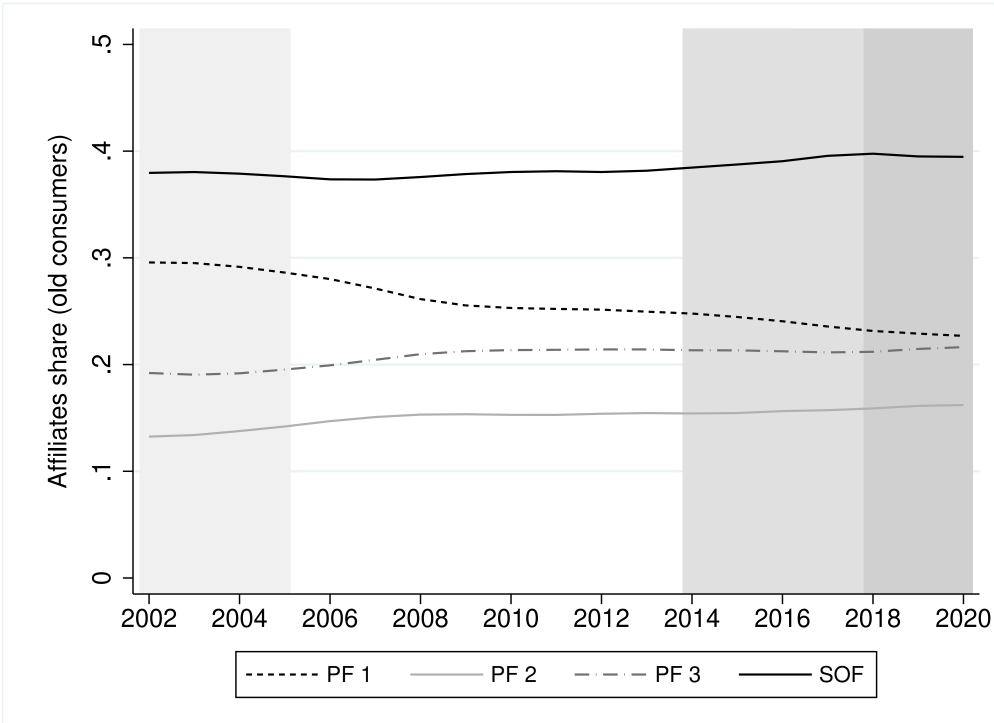
Notes. For each level of gross wages we display the average share of new enrollees by PFA. The three tertiles are composed of workers whose wages are below the compulsory enrollment threshold. The 4th group is composed of individuals above the threshold.

Table A.2: Share of sale force agents by PFA and period

Period	Firm			
	PF 1	PF 2	SOF	PF 3
'02-'05	0.21	0.24	0.35	0.21
'14-'17	0.22	0.15	0.36	0.27
'18-'19	0.24	0.15	0.36	0.25

Notes. Average share of the sales force by firm and period.

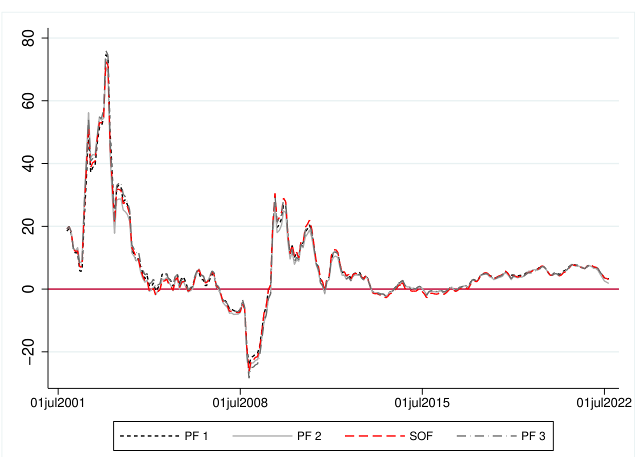
Figure A.1: Market share by PFA over time, all enrollees



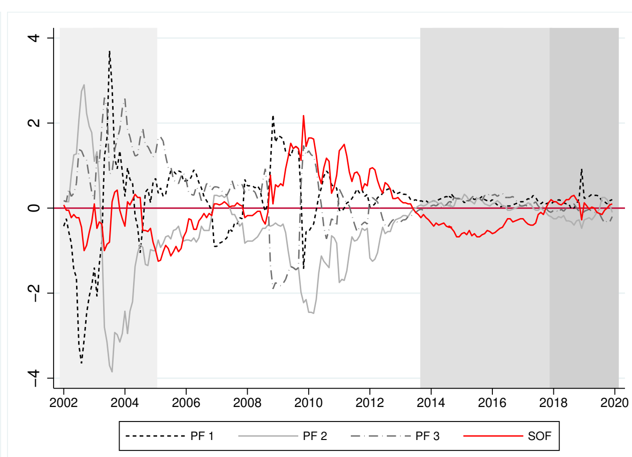
Note. Evolution of markets shares of old consumers by firm.

Figure A.2: Evolution of the real annual rate of return by firm

(a) Rate of return

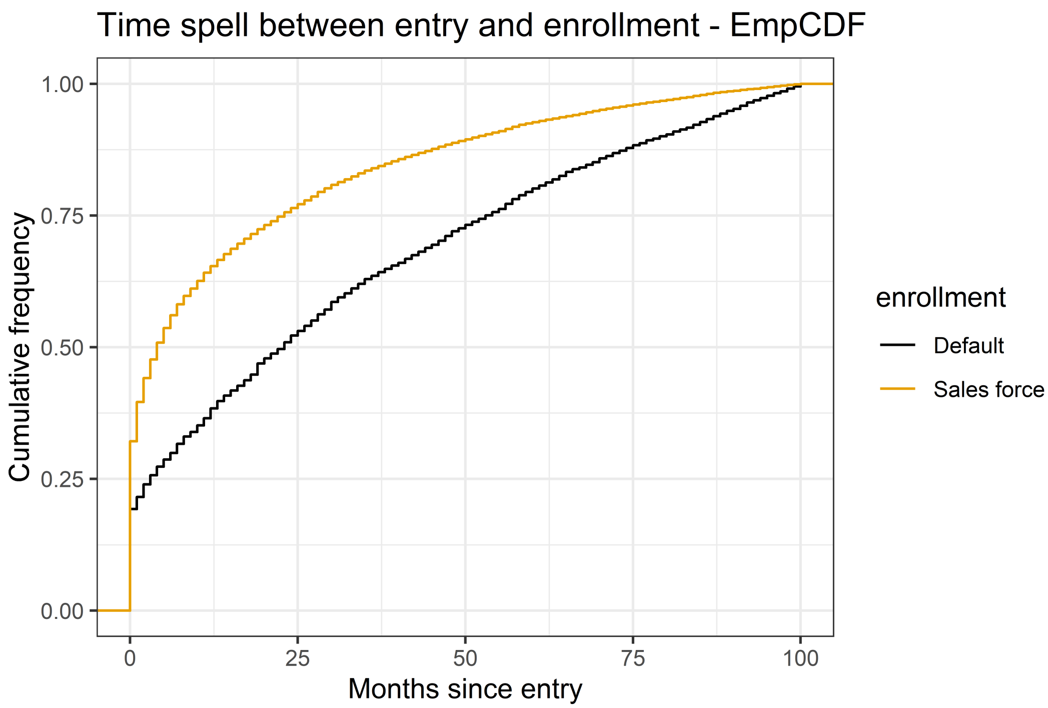


(b) Variation from the mean



Note. On the left, is the evolution of the real gross annual return in adjustable units, and on the right, is the variation with respect to the average.

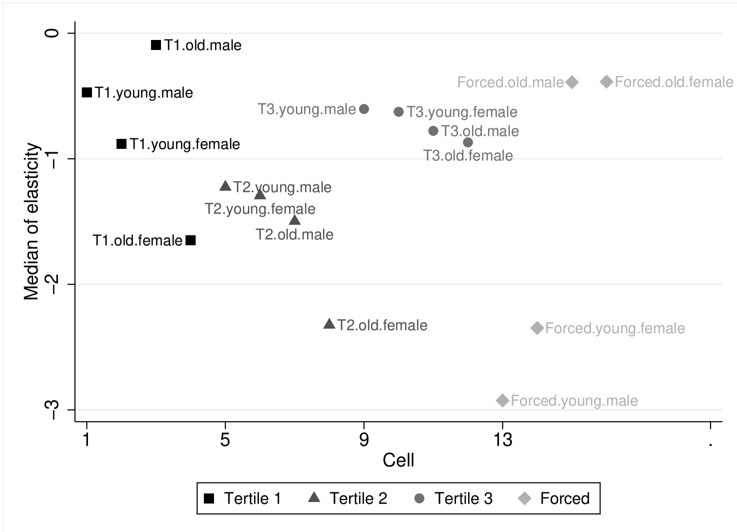
Figure A.4: ENTRY AND ENROLLMENT



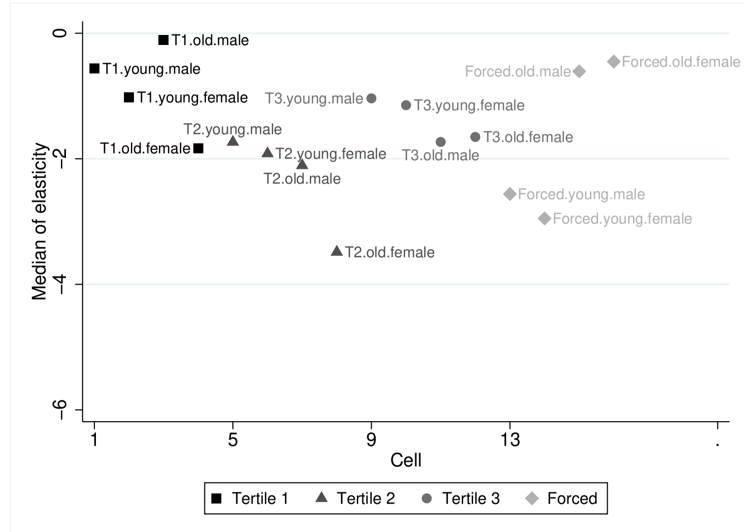
Note. Yellow curve: empirical cumulative distribution function of time spell between entry to the labor market and enrollment for individuals voluntarily affiliated. Black curve: cumulative distribution function of time spell between entry to the labor market and enrollment for individuals enrolled by default.

Figure A.5: MEDIAN FIRM ELASTICITY BY BRACKET AND PERIOD.

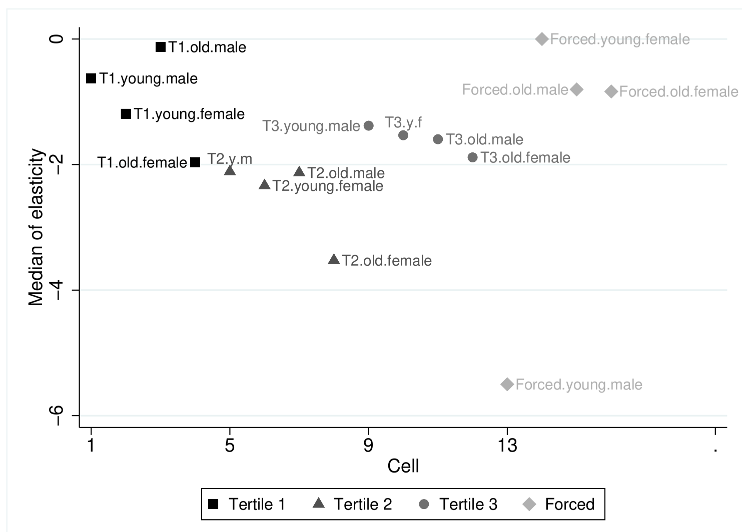
(a) 2002-2005



(b) 2014-2017



(c) 2018-2019



Note. Median elasticity by bracket and period. Elasticities are calculated at the observed fee levels and individual characteristics. Computed using estimates from equation (7) to generate the logit choice probability for each individual for each firm.

Table A.3: DEMAND ESTIMATION. CELLS 1 TO 8

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Total cost	-0.0073***	-0.0133***	-0.0032***	-0.0180***	-0.0147***	-0.0169***	-0.0138***	-0.0223***
	(0.0008)	(0.0009)	(0.0011)	(0.0011)	(0.0008)	(0.0009)	(0.0008)	(0.0010)
Cost Wage	3.0e-07***	5.3e-07***	1.9e-07***	6.7e-07***	3.4e-07***	4.0e-07***	3.4e-07***	5.2e-07***
	(4.3e-08)	(4.8e-08)	(5.6e-08)	(5.2e-08)	(2.7e-08)	(3.3e-08)	(2.6e-08)	(3.1e-08)
Tertile	1	1	1	1	2	2	2	2
Age	Young	Young	Old	Old	Young	Young	Old	Old
Gender	Male	Female	Male	Female	Male	Female	Male	Female
Outside option	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	36,975	35,390	30,195	38,395	47,690	39,045	27,150	27,070

Notes. Standard errors in parentheses*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Estimation of equation (7) via maximum likelihood. All specifications include fixed effects by firm and year interacted. Individuals that entered the market between 1997-2019. Cells 1 to 12 include an outside option and cells 13 to 16 no.

Table A.4: DEMAND ESTIMATION. CELLS 9 TO 16

	(9)	(10)	(11)	(12)	(13)	(14)	(15)	(16)
Total cost	-0.0034***	-0.0044***	-0.0029***	-0.0039***	-0.0051***	-0.0017	-0.00048*	-0.00059
	(0.0002)	(0.0005)	(9.45e-05)	(0.0002)	(0.0009)	(0.0017)	(0.0002)	(0.0005)
Cost wage	5.7e-09***	2.1e-08***	9.3e-10***	7.1e-09***	8.0e-09	-2.8e-08*	3.3e-10**	1.4e-09*
	(8.8e-10)	(6.4e-09)	(5.2e-11)	(1.0e-09)	(6.1e-09)	(1.6e-08)	(1.4e-10)	(8.2e-10)
Tertile	3	3	3	3	-	-	-	-
Age	Young	Young	Old	Old	Young	Young	Old	Old
Gender	Male	Female	Male	Female	Male	Female	Male	Female
Outside option	Yes	Yes	Yes	Yes	No	No	No	No
Observations	43,600	25,675	41,045	30,635	5,032	2,856	4,128	3,068

Notes. Standard errors in parentheses*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Estimation of equation (7) via maximum likelihood. All specifications include fixed effects by firm and year interacted. Individuals that entered the market between 1997-2019. Cells 1 to 12 include an outside option and cells 13 to 16 no.

Table A.5: SOF PLAYING AS A PROFIT-MAXIMIZING FIRM. FACTUAL 2014-2017.

	Fee (%)				Av. Δ	Outside	Profit		CW	ΔCW	TW	ΔTW
	PF 1	PF 2	PF 3	SOF	PF fee	option	PF	SOF		(%)		(%)
1- Factual 14-17	1.98	2.02	1.85	0.80	-	21	41	8	551	-	600	-
2- SOF NB	2.12	2.36	2.15	5.75	13	28	50	184	499	-9	734	22

Notes. Fee expressed as a % of the relevant gross wage for contribution to the sub-system. For-profits parameter $\lambda = 1$. Share old workers distributed equally between PFAs. CS consumers-workers welfare. TW total welfare. CS, Profit, and TW expressed in US$ millions per year.

Table A.6: COUNTERFACTUALS IMPLEMENTATION

		Marginal cost		For profit	
	Demand	PF	SOF	motive SOF (λ)	Share old
1- Factual 14-17	14-17	Backed out 14-17	Av. PF	Backed out 14-17	Observed
2- 4 PF	14-17	Backed out 14-17	Av. PF	-	Splitted equally
3- Optimal 14-17	14-17	Backed out 14-17	Av. PF	0	Observed
4- Factual 20	18-19	Backed out 14-17	Av. PF	Backed out 14-17	Observed
5- Optimal 20	18-19	Backed out 14-17	Av. PF	0	Observed
6- No caps 20	18-19	Backed out 14-17	Av. PF	Backed out 14-17	Observed

Notes. For the period 2020, we use the demand primitives of 2018-2019. For the same period, we use the backed-out MC of 2014-2017 because in 2018-2020 there is a cap on fees operating, so it is not possible to use the FOC to back-out the MC.

B Appendix: Salesforce average variable payments per enrollee

We define the wage mass of the sale force of firm j in month t as:

$$W_{jt} = L_{jt}^{sf}(w_{jt}^{fixed} + Aff_{jt} \times w_{jt}^{variable}) \tag{10}$$

being L_{jt}^{sf} the sales force, w_{jt}^{fixed} the average fixed wage, Aff_{jt} the average number of affiliates through sales force and $w_{jt}^{variable}$ the average variable wage.

We observe L_{jt}^{sf}, W_{jt} and Aff_{jt}. Regarding w_{jt}^{fixed}, we construct the following proxy. In Uruguay, the minimum wage by sector is set through collective bargaining and there is a specific negotiation group for PFAs' employees (group 11, Sub-group 1.3). The negotiation does not set the variable salary. However, the minimum wage set in the negotiation for the sales force (US$ 513 dollars) is taken as a proxy for the fixed-wage component, so the results obtained are an upper bound of the variable wage component.

C Appendix: Equilibrium fees

This section gives additional details of the model and the computation of the equilibrium fees. We begin by formalizing the Nash-Bertrand game on fees. In the next steps, we detail the PFA j's expected profits as a function of fees, costs, and demand primitives, starting from the individual level (i), used in the demand estimation, to reach the aggregate level that is used in equation 3 of the supply model.

C.1 Per-period profit function

C.1.1 Revenues

The per-period revenues function of firm j in period t is:

$$R_{jt} = f_j \times \Bigg(\underbrace{\sum_{i_n} w_{it}^n \times prob_{ij}(\mathbf{f}) \times (1-\rho_{it}^n)}_{\text{New workers}} + \underbrace{\sum_{i_o} w_{it}^o \times 1(d_i = j) \times (1-\rho_{it}^o)}_{\text{Old workers}} \Bigg) \tag{11}$$

f_j represents the fee of firm j, w_{it}^n and w_{it}^o the wage of the individual, $prob_{ij}$ the probability for individual i of choosing firm j which for new consumers depends on the vector of fees, and is the primitive recovered from the logit demand estimation:

$$prob_{ij}(\mathbf{f}) \quad = \quad \frac{\exp[\theta_{it} \times C_{ijt}(y_{it}, f_{jt}) + \delta_{jt}]}{1 + \sum_{k \in J} \exp[\theta_{it} \times C_{ikt}(y_{it}, f_{kt}) + \delta_{kt}]} \tag{12}$$

The old consumers already choose a firm, and they have infinite switching cost, so $1(d_i = j)$ takes the value 1 if the individual i_o choose firm j and 0 in other cases. ρ_{it} is the probability of retirement.

C.1.2 Costs

C_{jt} are the costs for firm j in period t:

$$C_{jt} = \underbrace{\sum_{i_n} prob_{ij}(\mathbf{f}) \times w_{it}^{n} \times \text{MC}_j}_{\text{Enrollment Cost of New Workers}} + \underbrace{F_j}_{\text{Fixed cost}} \tag{13}$$

MC_j is the marginal cost per dollar of wage and F_j firms' fixed costs. In relation to the fixed cost, the only regulatory requirement related to fixed costs is a minimal physical network with customer service offices in at least 5 departments (regions).

C.2 Net Present Value profits function

Now we discuss the net present value of the profits function, assuming that firms are committed to a single price in all periods from t to T. In this section, we start with the individual-level net present value equation for revenues and specify a series of assumptions made to transition to working at an aggregate level.

Assumption 1: wages w_{it} can be expressed as $w_{it} = w_i \times \omega_t$, where ω_t is the same for every person in period t. Therefore, the wage evolution curve over time for all individuals within a cohort is the same.

Assumption 2: Each year, a cohort of equal size, preferences, and salaries enters.

The net present value of revenues of firm j in period t expressed as a summation of cohorts is:

$$
\begin{aligned}
NPVR_{jt} \quad = \quad & f_j \times \Bigg(\underbrace{\sum_{i_n} prob_{ij}(\mathbf{f}) \times w_i^n \times \sum_{t=1}^{T} \beta^{t-1} \times \omega_t \times (1-\rho_{it}^n)^t}_{\text{New workers in 1, from 1 to T}} \\
+ \quad & \underbrace{\sum_{i_n} prob_{ij}(\mathbf{f}) \times w_i^n \times \sum_{t=2}^{T} \beta^{t-2} \times \omega_t \times (1-\rho_{it}^n)^{t-1}}_{\text{New workers in 2, from 2 to T}} + ... \\
+ \quad & \underbrace{\sum_{i_n} prob_{ij}(\mathbf{f}) \times w_i^n \times \sum_{t=T}^{T} \beta^{t-T} \times \omega_t \times (1-\rho_{it}^n)^{T-1}}_{\text{New workers in T, from T to T}} \\
+ \quad & \underbrace{\sum_{i_o} 1(d_i = j) \times w_i^o \times \sum_{t=1}^{T} \beta^{t-1} \times \omega_t \times (1-\rho_{it}^o)^t}_{\text{Old workers from 1 to T}} \Bigg)
\end{aligned} \tag{14}
$$

Assumption 3: The probability of retirement in each period is the same for all individuals within a cohort, so we can write $\rho_{it}^n = \rho_t^n$ and $\rho_{it}^o = \rho_t^o$. Therefore the third term of each sum does not depend on i and we call the terms as in $\sum_{t=1}^{T} \beta^{t-1} \times \omega_t \times (1-\rho_t^n)^t$ as $\gamma's$:

$$
\begin{aligned}
NPVR_{jt} \quad = \quad & f_j \times \Bigg(\underbrace{\sum_{i_n} prob_{ij}(\mathbf{f}) \times w_i^n \times \gamma_1}_{\text{New workers in 1, from 1 to T}} + \underbrace{\sum_{i_n} prob_{ij}(\mathbf{f}) \times w_i^n \times \gamma_2}_{\text{New workers in 2, from 2 to T}} + ... \\
+ \quad & \underbrace{\sum_{i_n} prob_{ij}(\mathbf{f}) \times w_i^n \times \gamma_3}_{\text{New workers in T, from T to T}} + \underbrace{\sum_{i_o} 1(d_i = j) \times w_i^o \times \gamma_o}_{\text{Old workers from 1 to T}} \Bigg)
\end{aligned} \tag{15}
$$

We denote $Z_n = (\gamma_1 + ... + \gamma_T)$ and $Z_o = \gamma_o$:

$$
NPVR_{jt} \quad = \quad f_j \times \left(\sum_{i_n} prob_{ij}(\mathbf{f}) \times w_i^n \times Z_n + \sum_{i_o} 1(d_i = j) \times w_i^o \times Z_o \right) \tag{16}
$$

Now call M the total mass of gross wages affected to the individual capitalization sub-system in period t, and α the share of that mass of wage that comes from new cohort. So $\sum_{i_n} w_i^n = \alpha M$ and $\sum_{i_o} w_i^o = (1-\alpha)M$.

We multiply and divide every term by the mass of wage of new and old, and call $s_j^n = \sum_{i_n} \frac{prob_{ij}(\mathbf{f}) \times w_{it}^n}{\sum_{i_n} w_i^n}$ and $s_j^o = \sum_{i_o} \frac{prob_{ij} \times w_{it}^o}{\sum_{i_o} w_i^o}$, so we get:

$$NPVR_{jt} = f_j \times \left(s_j^n \times \alpha \times M \times Z_n + s_j^o \times (1-\alpha) \times M \times Z_o \right) \tag{17}$$

This expression is equivalent to the term associated with revenues from equation 3. Finally, the net present value of firm j in period t is:

$$V_{jt} = NPVR_{jt} - \frac{(1-\beta^T) \times MC_j \times \alpha \times M_t \times s_j^n(\mathbf{f})}{(1-\beta)} - \frac{(1-\beta^T) \times F_j}{(1-\beta)} \tag{18}$$

We assume the existence of a pure-strategy Nash-Bertrand equilibrium in fees, and that the fees that support it are strictly positive (and lower than 100% of the contribution). In this context, a Nash-Bertrand equilibrium in this game is a vector of fees (f_j), such that

$$(f_j) \in arg\, max_{\{f_{jt}\}_0^T} V_{jt}(\mathbf{f}|W^o, W^n, MC_j) \tag{19}$$

for each PFA $j \in J$.

Regarding the way of computing the equilibria, empirically we have a relevant difference with Hastings et al. (2017), where from the demand estimation it emerges that a substantial number of individuals have price elasticities of demand that are positive or close to zero. That creates problems for calculating equilibrium because the maximization problem is not convex. In our case, although there are some inelastic demand segments, there are no individuals with positive elasticities, which a priori allows us to use the standard solution to a logit-Bertrand price game. Therefore, to solve the counterfactuals, we compute the 4 first-order conditions of the net present value of the profits of the firms with respect to fees, as expressed in equations 8 and 10, with the function fmincon in matlab. There are no convergence problems. To address the multiplicity of equilibria, we use random starting values to calculate the equilibrium fees.

Mark ups and pass-through in small and medium retailers for rice, tomato sauce and oil

Pablo Blanchard *

Abstract

In this paper, we recover and decompose markups, and estimate the pass-through rates from cost to prices in small and medium retail stores for oil, tomato sauce, and rice in Uruguay using a structural model of demand and assumptions about the competitive behavior of producers. The market power for these products has been under the study of the Commission of Promotion and Defence of Competence since 2016, and the proposed methodology allows for deepening in the measure and the understanding of the origin of that market power. The markups for oil and tomato sauce are around 25% for Nash Bertrand competition assumption, and 50% for the collusion assumption, while for rice are 36% and 75% respectively. For its part, about 65% of the market power under Nash Bertrand assumption is explained by the ability of firms to differentiate products and 35% for the ownership structure in the case of oil and sauce. In the case of rice, 49% are explained for differentiation and 51% for ownership structure. Finally, the pass-through rates are low for the three products, being under both behavioral assumptions lower than 55% for the three products.

JEL Classification: D43, L11, L81

Key words: Market Power, Cost Pass-Through, Discrete choice models, Product differentiation

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1 Introduction

In the last years, the formation of prices for articles of the basic basket has been in the public debate in Uruguay and there have been incipient efforts by the Commission for the Defense and promotion of Competition (CDPC) to analyze the existence of anti-competitive practices in these markets. In particular, in 2016 the commission carried out an analysis of price formation for a series of specific products (oil, rice, tomato sauce, and bread), on which it wanted to know the market power exercised by firms and the possible existence of anti-competitive practices in price formation.[1]

As a result of this analysis, implemented from descriptive statistics on quantities sold, consumer prices, and prices paid by the supermarket chain to the producers, the margins of the main supermarket chains for these products are known, as well as a general characterization of concentration in each market (Czarnievicz and Zipitría (2018)). However, as the report itself points out, with this methodology it is difficult to deepen the analysis.

This paper seeks to delve into the analysis of price formation, market power, and pass-through from cost to price for 3 of the aforementioned items: oil, rice, and tomato sauce. In particular, the markups and the capacity of the production firms to transfer cost shocks to prices (pass-through) consistent with the demand information and different assumptions about the way in which the production firms compete are estimated. For its part, it breaks down market power into two sources: generated by product differentiation and generated by the fact that firms offer different varieties of the same product (portfolio effect).

The analysis focuses on producers and there is information on quantities and sales prices for small and medium-sized retail stores from April 2016 to January 2018. We estimate a structural model of demand for differentiated products according to the methodology proposed by Berry et al. (1995) (BLP approach), using information on prices and quantities, relieved from scanning at the time of purchase. The demand estimate is then used in conjunction with different assumptions about competition among firms to recover marginal costs and markups. Following the methodology of Nevo (2001), the different estimated marginal costs and markups are compared with each other and with descriptive statistics of production costs from other sources of information to separate between different sources that explain the observed margins. Finally, it is analyzed how different simulated cost shocks are transferred to sales prices by firms.

In terms of the relevance of the problem addressed, measuring the capacity of firms to increase their price above the marginal cost of production is a central problem in the literature on industrial

[1] Resolution No. 31/016: https://www.gub.uy/ministerio-economia-finanzas/institucional/normativa/resolucion-n-31016-asunto-12-2016-estudio-preparatorio-mercados-productos

organization and is usually the measure used to analyze the market power that a firm can exercise. For its part, the pass-through from costs to prices is an issue of great importance for understanding price formation at the microeconomic level and developing competition defense policies. Understanding how firms transfer increases in their costs to prices and how these increases depend both on the behavior of firms and demand is a relevant problem in economics.

On the one hand, the article is related to the empirical literature that focuses on the analysis of firms' market power, usually understood as the ability of firms to increase prices above the marginal costs. This is also a key concern for competition policy, not only in terms of the measure itself but also what is the origin of this market power (Bet (2021)). As market power is a function of demand and cost primitives, but also of the firm conduct, its estimation presents several difficulties. Some literature has focused on how to distinguish between competing oligopoly models based on an IV approach (e.g, Bresnahan (1982), Berry and Haile (2014), Duarte et al. (2021)). As is pointed out in Bet (2021), identifying firm conduct following this approach is difficult because of the nature of the required instruments, which often are weak. There is also the production function approach (Hall (1988), De Loecker and Warzynski (2012), De Loecker et al. (2016), De Loecker and Scott (2016), Raval (2020), Doraszelski and Jaumandreu (2019), among others) which relies on production data, and some recent efforts to use simultaneously both approaches (Bet (2021), De Loecker and Scott (2016)). Finally, Nevo (2001) or Slade (2004) use information on production cost from another source and compare it with the results obtained from the demand estimation and observed prices, and different competing oligopoly models. This last alternative is followed in this article due to the fact that the Uruguayan market is very concentrated and there are few firms producing each item, so it is not possible to reliably apply the production function approach.

On the other hand, this article is related to a strand of empirical and theoretical literature that addresses the pass-through from costs to prices from the industrial organization literature. As pointed out in Kim and Cotterill (2008), the first theoretical antecedents of this literature focus on the cases of perfect competition and monopoly. In the first case, the pass-through from costs to prices is determined by the elasticities of supply and demand, being greater the more inelastic the demand and the more elastic the supply, reaching one hundred percent when the supply is infinitely elastic. For its part, in the case of a monopoly, the pass-through of costs to prices depends on the slope of the demand curve and the elasticity of the cost function to changes in quantity. Bulow and Pfleiderer (1983) show that, if marginal cost is constant and demand is linear, the cost-to-price pass-through coefficient for a monopoly is fifty percent. In Weyl and Fabinger (2013) it is shown that, under different assumptions about the cost function and the curvature of the demand function, the pass-through of costs to prices in monopoly is not bounded, and may be greater than one hundred or less than fifty percent.

As a way to overcome the limitations implied by assuming perfect competition or monopoly, several studies have theoretically analyzed the pass-through from costs to prices in the presence of imperfect competition (Stern (1987), Katz and Rosen (1983), Delipalla and Keen (1992)), but focusing on homogeneous products and competition in quantities.

Among the most relevant theoretical antecedents of works that have focused on the pass-through of costs to prices in markets with differentiated products are the works of Anderson et al. (2001), where the incidence of taxes in an oligopolistic industry is analyzed when competing for prices in differentiated markets and Froeb et al. (2005), where the effect of company mergers on prices is analyzed.

On the other hand, there are theoretical antecedents from the marketing literature, among which Moorthy (2005) stands out, who proposes a comparative static analysis of the pass-through from costs to prices for retailers, when there are multiple retailers competing and each one sells multiple varieties of a product. By incorporating multiple varieties of each product, the dimension of the retailer is introduced as a manager who is interested in the joint profits derived from the sales of the different varieties of a product. The paper focuses on how different types of cost changes (distinguishing for example between aggregate, brand-specific, and store-specific shocks), as well as inter-retailer and inter-brand competition, affect the pass-through from costs to prices.

There is also literature that empirically addresses this problem, mostly through reduced-form analyses with industry-level data (Sullivan (1985); Karp and Perloff (1987); Besley and Rosen (1999)). Regarding the empirical antecedents that use structural models, two works stand out for their proximity to the methodology that will be used. In the first place, the work of Nevo (2001), in which market power in the cereal industry for the United States is studied, estimating a demand system for differentiated goods at the brand level and using the estimated parameters together with assumptions about competition between producers to recover marginal costs and profit margins. Second, the Kim and Cotterill (2008) study, which constitutes the empirical background most directly related to the present work. The authors study the pass-through of costs to prices in the processed cheese market in the United States, estimating a structural model of demand for differentiated goods that they use to recover the marginal costs of each product under different supply assumptions (collusion or Bertrand prices competition) and then calculate the pass-through from costs to prices. In the aforementioned work, the results obtained by structural estimation are compared with those obtained based on reduced-form equations. From this, they conclude that the processed cheese market in the United States operates in imperfect competition, with a level of competition greater than absolute collusion, but less than full competition in prices.

The empirical literature on the pass-through of costs to prices has focused on the study of markets in developed countries, and no precedents have been found for developing countries. The study of the pass-through of costs to prices in consumer goods is potentially different in developing economies due to possible differences in the levels of existing competition, as well as in the regulation for its promotion. In economies with less intensity of competition, it is expected to find higher price levels and markups, which may cause cost shocks to be transmitted to final prices to a lesser extent. To the best of our knowledge, this phenomenon has not been studied empirically for consumer goods, and one of the limitations is given by the scarcity of markets for consumer goods with low levels of competition that are registered in developed countries. Along these lines, Mahoney and Weyl (2017) theoretically discusses the possibility that lower levels of competition are characterized by the lower transmission of cost shocks to prices in markets with adverse selection.

At the national level, there are few studies that analyze the microeconomic behavior of consumers and suppliers and its link with prices and margins. Among them, Borraz et al. (2012) studies the setting of prices at the supermarket level, finding that they change approximately 5 times a year, without a seasonal pattern, with a high synchronization and concentration on the first day of the month. In Aguirre et al. (2022) the price differential between the retailers authorized to sell to beneficiaries of a cash transfer program (also medium and small size retailers) regarding supermarket chains is studied. Beneficiaries are found to pay prices significantly higher than what they would get in nearby larger stores that do not participate in the program. For their part, in Rius and Zipitría (2016) a typology of prices is carried out according to their formation mechanism, as well as an analysis of the prices of products in supermarket chains, in which no price trend is found above the inflation. Despite this, products and establishments with the potential to set prices above what they would set in competitive markets are found. Finally, Borraz et al. (2016) study the effects of political borders on relative prices between regions, finding that previous studies on the subject systematically underestimated transportation costs, and proposing an alternative methodology that corrects the bias.

The next sections are organized as follows. Section 2 introduces a short description of the setting. In Section 3 we describe the data, present detailed information about how we work with it, and statistics descriptives of the industry. Section 4 describes the model, the estimation procedure, and identification. Results are presented in Section 5 and finally, Section 6 concludes.

2 Setting

Our focus is on the estimation of demand and recovering markups and pass-through for rice, oil, and tomato sauce during the period from April 2016 to January 2018. The three products analyzed have in common the fact that they are mainly produced in Uruguay and the domestic market shows oligopolistic structures with a concentration in one or two firms with the majority of the market share. From Czarnievicz and Zipitría (2018) we know that during the period 2014-2016 and in supermarket chains, for oil the firm COUSA represented approximately 69% of the market share, for rice there was a duopoly between SAMAN (41%) and Coopar (47%) and in tomato sauce Barraca Deambrosi presented a clear leadership with 47% of the market share, followed by Conaprole (21%).

The rice sector has very different characteristics from the others, since Uruguay is a world-class rice exporter, with high productivity in its production. Since 1959, the price paid by the mills to the producer has been determined between the Rice Growers Association and the Rice Mills Guild, based on production costs, the domestic price, and the export price. In 2017 the estimated consumption of white rice in the domestic market was 45.000 tons[2], which represented approximately 5% of what was exported (Miraballes Iguiní (2021)), so we are studying a market that represents a small fraction of the total sales of this firms.

Tomato production reached 39,000 tons in the 2016/17 harvest, while the imported volume was 1,826 tons (Yearbook 2017, OPYPA). Nevertheless, of the total tomato produced, 20% is destined for processing. Therefore, national production covers between 20 and 25% of industrialized tomato consumption and the rest of the raw material is imported.[3] Regarding the tomato processing industries, in a 2014 resolution, the Commission for the Promotion and Defense of Competition sanctions 5 companies and exonerates 1 for carrying out an illicit agreement for anti-competitive purposes (Resolution No. 24/014).

Regarding the oil, in the present work sunflower, soybean and canola oil are included as different varieties of the same product. In 2017 the estimated domestic consumption of these varieties of oil was 74.000 tons[4]. During the analyzed period, the agricultural production of oil in Uruguay shows a significant increase in soybean production, accompanying the regional evolution and a pronounced decrease in sunflower production. For its part, the oilseed industrial sector is made up of a single company (COUSA), which in 2015 had an installed capacity of approximately 1,450 tons per day. The oil sector in Uruguay is strongly protected against imports (Brum et al. (2012)). During the period analyzed, the tariff rate for imports from Argentina was 16% (Ordinance No. 643/006) and for areas outside Mercosur it was 21%. Nevertheless, there is some degree of imported products and the competitors are mainly products imported from Argentina (Horta et al. (2017)).

In the present study, the margins and the pass-through of costs to the final prices of the products

[2]https://www.indexmundi.com/agriculture/?pais=uyproducto=arroz-blancovariable=consumo-domesticol=es
[3]INIA report http://www.ainfo.inia.uy/digital/bitstream/item/4878/1/hd-105-tomate-industria-Oct.2011.pdf
[4]https://www.indexmundi.com/agriculture/?pais=uyproducto=aceite-de-semilla-de-girasolvariable=consumo-domesticol=es

are analyzed. Nevertheless, the available data do not allow an analysis of the vertical links between producers, distributors, and retailers. Similar to Nevo (2001) and Kim and Cotterill (2008), the focus is on the competition between producers. We assume that retail margin is an additional cost to producers, which is consistent with a wide variety of models of manufacturer-retailer interaction. Additionally, following standard practice in the BLP approach, we treat the retail industry as a price-taking, perfectly competitive industry, which implies that the store and product-level elasticities are identical. This is a problematic assumption, particularly when working with large supermarket chains. As mentioned in the description of the data, in this article we are working with small and medium retailer firms, but still, we should be cautious with this simplification.

As these are oligopolistic markets in which there are few producers, it is not possible to make good estimates with the production function approach. As an imperfect approximation, descriptive data are presented in Table 1, containing information on production costs, collected from the Economic Activity Survey (EAE) 2016, for companies in whose total sales, the analyzed products weigh at least 50%[5]. The gross price-average variable cost margin for these industries is 30.1% for oil, 34.6% for tomato sauce and 39.7% for rice[6] It is relevant to emphasize that what we recover in the Table is the gross margin and not the markup. To recover the mark up we need the marginal costs, which we cannot obtain from descriptive information from surveys.

Table 1: Aggregate descriptive of production costs

	Oil		Tomato sauce		Rice	
	Mill pesos	% of value	Mill pesos	% of value	Mill pesos	% of value
Sales	6531	100	3530	100	23919	100
Materials	3890	59.6	1765	50	15565	65.1
Labor	607	9.3	525	14.9	2027	8.5
Energy	70	1.1	18	0.5	417	1.7
Gross margin (GM)	1964	30.1	1222	34.6	5910	24.7
GM internal market	-	-	-	-		39.7

Notes. Source: Economic Activity Survey (EAE, 2016). Sector: four-digit ISIC revision 3. we use firms in which at least 50% of sales correspond to the analyzed product.

[5] In this Survey only firms with at least 50 employees are mandatorily included in the sample.

[6] In the case of rice, we can decompose the gross margin in the internal and external market, knowing the sales (PxQ) to the internal and external market (EAE), quantities exported and export prices (MGAP - DIEA. FLAR) and sales prices of the producers to the big national retailers (Czarnievicz and Zipitría (2018)), assuming that the margin is the same for small retailers (in the reality the margin in the internal market is a weighted average between the margins obtained from the big chains and medium/small retailers).

3 Data and descriptives

A novel database is used that consists of a panel with prices and quantities sold in each store for different varieties of the three products studied, defined at the UPC level, for the main localities of Uruguay, with daily information for the period between April 2016 to January 2018. The information comes from the scan at the time of purchase and is provided by a Point of Sales (POS) provider, which specializes in providing this service to small and medium-sized retailers, and is one of the leading firms in this segment of retailers. We observe the universe of sales in the retailers who operate with this provider.

In order to have a notion of how much the sales in these stores represent in the total consumption of the studied products, Table 2 shows the percentage represented by the quantities sold during 2017 in the available database in relation to the total consumption of white rice and oil (sunflower plus soybean) in Uruguay, reported in Index Mundi. The database represents 18% of the total consumption of rice and 12% of the total consumption of oil in 2017 in Uruguay.

Table 2: Sales in sample and consumption in Uruguay in 2017

Product	Total sales (quantities)	Consumption in Uruguay	Ratio
Oil (soy + sunflower)	8.89	74	12%
Rice	8.15	45	18%

Notes. Total sales are the quantities of kilos or liters sold in all available stores in the database during 2017. Consumption in Uruguay is the total consumption of the goods reported in index mundi for 2017. Own elaboration based on the database and https://www.indexmundi.com/. There is no information about tomato sauce.

As can be seen in Table 3, in general terms we observe small and medium-sized retailer firms from 50 different regions of Uruguay. The median retailer in the database sells 127 kilos (or liters) of the product by month. The main chains of supermarkets, analyzed in Czarnievicz and Zipitría (2018) are not included. One limitation of the database is that there is no additional information on the size of the stores (beyond the total sales and quantities of these products) or whether they belong to a chain.

Given the focus of the article and based on the main related literature, a series of decisions are made on how to work with the data. The locality and not the store is taken as the unit of analysis, taking the aggregate of sales in the locality and the average prices (weighted by the number of sales in each store). This is because when dealing with small stores, we want to avoid the presence of zeros in the base, which are problematic for the BLP estimate. For its part, since the interest lies in the competition between producers, it is understood that it is a reasonable simplification. A third reason is that, since we do not have georeferenced information on the location of the stores, to work at the store level we should assume that they are local monopolists (Chidmi and Lopez (2007)). In relation to the period of time, we work at the quarter level (we observe 8 quarters), also in line with the main related literature. Regarding the different presentations of each product (for example 900 milliliters or 500 milliliters oil), all presentations are expressed as a price per liter or kilo depending on the item.

Additionally, to rule out varieties of products whose sales levels are insignificant at the national level, only varieties with at least a 2% market share at the national level are taken.

Table 4 presents descriptive of the demand information available for the three products, including the producer, the variety, the average price and share for the entire period, and the main observable characteristics (this information is complemented with Figures A.1, A.2 and A.3 of the appendix, with the evolution of shares and prices for each variety).[7] For oil, we observe 9 varieties and 5 different producers. COUSA is the leading firm, with 78% of the market (in line with Czarnievicz and Zipitría (2018)) and 5 varieties. There is a clear distinction in terms of prices between soy oils and the rest, being the first one the cheaper for the entire period. Nevertheless, it is observed a sharp reduction in the prices of sunflower and canola oil mainly produced during 2016. It is also important to note that while there is competence in the segment of soy oil, that is not the case in the canola and sunflower segments.

For tomato sauce, we observe 10 varieties produced by 6 firms. The leading firm is Barraca Deambrosi with 57% of the market, which produces 4 varieties. In this market, there are competitors in different segments of quality and price. Conaprole offers a concentrated variety with the highest price, while Don Perita offers a variety with the lowest price. The evolution of prices and shares is relatively stable in the period, with an increase in the prices of Conaprole and a decrease in the prices of Cololo. For its part, Big presents a slight decrease in prices and a marked increase in share.

In the rice market, we observe 10 varieties produced by 4 firms, and a duopoly of Saman and Coopar, with 38% and 57% of the market, producing 4 varieties each one. In Figure A.3 we can see a clear pattern in terms of prices and observable characteristics: varieties of type one (at least 90% of entire grains) with prices approximately between 35 and 40 Uruguayan pesos and varieties of lower quality with prices between 23 and 28 Uruguayan pesos. While for varieties of lower quality, there is competition between the 4 firms, in the high-quality segment there are only varieties of the two main firms.

In the last column of Table 4 we can observe the percentage of retailers in which each variety of the

[7]In the descriptives, we show current prices. In the main specifications, we deflate the prices using as base 2016. Nevertheless, the main results hold with current prices.

Table 3: Descriptive statistics of retailers by locality

Department	Locality	Municipality	Retailers	Quantities sold		
				p25	p50	p75
Artigas	Artigas		5	26	30	81
Artigas	Bella Union		4	159	343	653
Canelones	Barros Blancos		8	28	192	2904
Canelones	Canelones		8	292	1013	2615
Canelones	La Paz		5	96	519	1185
Canelones	Las Piedras		20	57	909	2893
Canelones	Pando		12	64	156	253
Canelones	Parque Del Plata		5	471	748	2047
Canelones	Paso Carrasco		7	60	902	2870
Canelones	Pinar		8	43	58	727
Canelones	Progreso		10	27	95	163
Canelones	Salinas		8	86	303	867
Canelones	San Ramon		6	49	671	845
Canelones	Santa Lucia		8	176	802	1293
Canelones	Sauce		7	49	125	215
Canelones	Solymar		9	379	992	1180
Canelones	Toledo		7	30	84	357
Cerro Largo	Melo		11	36	246	373
Colonia	Carmelo		7	139	360	3441
Colonia	Colonia		3	83	618	1307
Durazno	Durazno		6	104	705	3448
Flores	Trinidad		7	55	623	2116
Florida	Florida		15	159	664	2067
Lavalleja	Minas		6	588	1498	2981
Maldonado	Maldonado		31	224	760	1654
Maldonado	Piriapolis		5	33	352	1594
Maldonado	Punta Del Este		2	86	250	414
Maldonado	San Carlos		6	176	687	3574
Montevideo	Montevideo	A	49	64	261	687
Montevideo	Montevideo	B	72	32	125	425
Montevideo	Montevideo	C	34	129	218	736
Montevideo	Montevideo	CH	49	55	207	422
Montevideo	Montevideo	D	36	130	263	687
Montevideo	Montevideo	E	36	75	227	722
Montevideo	Montevideo	F	38	141	374	2227
Montevideo	Montevideo	G	33	74	419	1314
Paysandu	Paysandu		18	34	104	360
Rio Negro	Fray Bentos		9	70	707	882
Rio Negro	Young		7	55	207	1749
Rivera	Rivera		14	23	102	235
Rocha	Rocha		12	97	172	728
Salto	Salto		22	96	437	738
San Jose	Ciudad Del Plata		11	156	260	713
San Jose	Libertad		8	126	368	713
San Jose	San Jose De Mayo		17	33	292	1032
Soriano	Dolores		5	22	60	466
Soriano	Mercedes		4	80	172	2626
Tacuarembo	Paso De Los Toros		2	125	474	823
Tacuarembo	Tacuarembo		11	33	86	535
Treinta Y Tres	Treinta Y Tres		8	114	212	1765

Notes. Descriptive for retailers that register sales throughout the period. The last 3 columns show the average monthly quantities of units of products sold by retailers in the 25th, 50th, and 75th percentiles of each location. To build the percentiles, a pool of the 3 products is made, where the kilos for tomato sauce and rice and liters for oil are added.

Table 4: Descriptive statistics of products

Product	Producer	Variety	Price	Share	Characteristic			% of retailers selling
					Soy	Canola	High ol.	
Oil	COUSA	Optimo canola	64	4	0	1	0	52
Oil	COUSA	Optimo girasol	73	11	0	0	0	84
Oil	COUSA	Optimo girasol altoleico	85	2	0	0	1	49
Oil	COUSA	Uruguay girasol	60	16	0	0	0	71
Oil	COUSA	Condesa soja	44	47	1	0	0	86
Oil	De diez	De diez soja	45	6	1	0	0	6
Oil	Demas	Demas soja	47	5	1	0	0	32
Oil	Revelacion	Revelacion soja	45	4	1	0	0	32
Oil	Soldo	Rio de la plata soja	47	4	1	0	0	31
					Conc.			
Tomato sauce	Barraca Deamb.	Qualitas	33	4	0			22
Tomato sauce	Barraca Deamb.	Gourmet	50	22	1			72
Tomato sauce	Barraca Deamb.	De ley	46	28	0			62
Tomato sauce	Barraca Deamb.	Gourmet napolitana	51	3	1			37
Tomato sauce	Big	Big	37	12	1			34
Tomato sauce	Cololo	Cololo	49	3	0			35
Tomato sauce	Conaprole	Conaprole	59	12	1			81
Tomato sauce	Don Perita	Don Perita	30	5	0			32
Tomato sauce	Rigby	Rigby	33	6	0			28
Tomato sauce	Rigby	Rigby italiana	40	6	0			24
					Type 1	Parbo		
Rice	Casarone	Casarone	23	2	0	0		20
Rice	Coopar	Blue patna parboiled	39	2	1	1		34
Rice	Coopar	Shiva patna	23	19	0	0		49
Rice	Coopar	Blue patna	37	8	1	0		70
Rice	Coopar	Green chef	36	28	1	0		87
Rice	SAMAN	Saman patna	34	6	1	0		43
Rice	SAMAN	Saman parboiled	40	4	1	1		62
Rice	SAMAN	Saman	36	16	1	0		75
Rice	SAMAN	Aruba patna	24	12	0	0		51
Rice	San Jose	San jose	23	4	0	0		18

Notes. Varieties with at least 2% of the market share at the national level. Average prices and market shares for the entire period. Prices are expressed in Uruguayan pesos by liter for oil and sauce and by kilo for rice. Type 1 takes the value 1 if the rice variety has more than 90% of entire grains. Parboiled takes the value 1 if the rice variety is partially boiled. Soy and Canola refer to the plant from which the oil is extracted (the third type in the database is sunflower). Finally, high oleic oil is one that contains at least 75% oleic acid in its composition.

product is available. It is important to note that when we aggregate information at the locality level and calculate substitution patterns at that level, we are ignoring the fact that these different varieties could not be available in the same retailer.

Additionally, as demographic characteristics for the BLP model, we include information from the national census of 2011, related to age, education years, and sex by region. Descriptive statistics by region are reported in Table A.1 of the appendix.

Finally, a relevant problem in demand estimates is given by the determination of the size of the market for the calculation of market shares. As pointed out in Nevo (2000), the total size of the market must be defined according to the context and the particularities of the problem addressed but making sure that it is large enough to avoid the market shares of the external option is worth zero (the outside option includes the possibility of not acquiring the product or acquiring it in a store that is not part of the database). For example, Nevo (2001) in his study on the cereal market, assumes that the size of the market is one serving of cereal per capita per day, while Bresnahan et al. (1997) to estimate the demand for computers, take the number of office employees.

In the present study, the unit of analysis is the locality quarter. The market size is determined in relation to the number of inhabitants of each locality, taken from the 2011 census and scaled as follows: for tomato sauce and rice it is assumed that one liter/kilo per quarter is consumed, and for oil, it is assumed that two liters are consumed (maintaining the relationship between rice and oil consumption observed in Table 2)[8].

4 Model

The strategy consists of estimating the demand system for each product at the locality-quarter level, modeling it as a function of product characteristics and consumer preferences. That demand information is then used in conjunction with assumptions about how producers compete to estimate marginal costs, margins, and finally the pass-through from costs to prices consistent with estimated demand.

4.1 Demand

The demand system is estimated using a logit model of random coefficients. This type of model makes it possible to incorporate the heterogeneity of consumer preferences for the observed and unobserved characteristics of the products. We follow the methodology proposed by Berry et al. (1995) and Nevo (2001), which allows estimating this type of model with aggregate demand data (that is, without knowing what each individual buys).

It is assumed that with each purchase, consumers buy one unit of the product and choose the variety that offers them the greatest utility. Following Kim and Cotterill (2008), the indirect utility of consumer i, for variety j in market m is given by $U_{ijm}(x_{jm}, \xi_{jm}, p_{jm}, D_i, v_i; \theta)$, being x_{jm} the observed

[8]It is important to bear in mind that this market size should not represent total consumption, but potential purchases in these stores.

characteristics of the product, ξ_{ijm} unobserved characteristics of the product, p_{jm} prices, D_i observed characteristics of consumers, v_i characteristics unobserved of the consumers. For its part, θ is a vector of unknown parameters to be estimated.

The introduction of observed characteristics of consumers does not require information at the individual level, but rather it can be extracted from the empirical distribution of these characteristics in the population, as is done in the present work.

The indirect utility function is defined as:

$$u_{ijm} = \beta_i x_{jm} - \alpha_i p_{jm} + \xi_{jm} + \epsilon_{ijm} \tag{1}$$

Where α_i is the marginal utility of income of consumer i, β_i represents specific individual parameters related to other product characteristics different from price and ϵ_{ijm} is a zero-mean stochastic term.

Let be $\theta = \theta(\theta_1, \theta_2)$ a vector containing the parameters of the model. The vector $\theta_1 = (\alpha, \beta)$ contains the linear parameters, while $\theta_2 = (\Pi, \Sigma)$ contains the non-linear parameters.

Indirect utility can be divided into two parts:

$$u_{ijm} = \delta_{jm}(x_j, p_{jm}, \xi_{jm}; \theta_1) + \mu_{ijm}(x_j, p_{jm}, v_i, D_i; \theta_2) + \epsilon_{ijm} \tag{2}$$

$$\delta_{jm} = \beta x_{jm} - \alpha p_{jm} + \xi_{jm} \tag{3}$$

$$\mu_{ijm} = [-p_{jm}, x_j](\Pi D_i + \Sigma . v_i) \tag{4}$$

On the one hand, the average utility of variety j in the market m (δ_{jm}), and on the other, the deviation from the average utility, which captures the effect of the random coefficients (μ_{ijm}).

The deviations from the mean utility (μ_{ijm}) depend on the observed characteristics of the individuals D_i, and the unobserved characteristics v_i. The distribution of the parameters of consumer preferences by the characteristics of the products is modeled as:

$$\begin{pmatrix} \alpha_i \\ \beta_i \end{pmatrix} = \begin{pmatrix} \alpha \\ \beta \end{pmatrix} + \Pi D_i + \Sigma v_i \ , \quad v_i \sim P_v^*(v) \ and \ D_i \sim \hat{P}_D^*(D) \tag{5}$$

being $P_v^*(.)$ a parametric distribution and $\hat{P}_D^*(.)$ a non-parametric distribution extracted from the 2011 population census. Π is a (K+1)xd matrix with coefficients representing how tastes for product characteristics vary with demographic characteristics. The unobserved individual characteristics are taken from random draws from a multivariate normal distribution, that is $P_v^*(.)$, is a $N(0, I_{k+1})$. In this way, individual heterogeneity is introduced in the taste for the characteristics of the products. One draw is taken for each individual for each product characteristic used, plus one for price (hence the K+1). In this context, it is an array of parameters of dimension (K+1)x(K+1), which allows each component to have different variances and correlations between characteristics. It is assumed that v_i and D_i are independent.

The specification of the demand closes with equation (6), which represents the utility of an external option (outside option or outside good), which is normalized in such a way that $\mu i0m = \epsilon i0m$. If the outside option is not included, a simultaneous increase in the price of domestic goods does not result in any change in aggregate consumption. The market share of the outside option is defined as the total size of the market minus the sum of the market shares of the inside goods.

$$\mu_{i0m} = \xi_{0m} + \Pi_0 D_i + \omega_0 v_{i0m} + \epsilon_{i0m} \tag{6}$$

A_{jm} is defined as the set of values D, v and ϵ that induce the choice of j in market m. D, v, and ϵ are assumed to be independently distributed. D is taken from an empirical distribution F, obtained from the national census of 2011, and v is taken from a multivariate normal distribution N. For its part, it is assumed that ϵ has an extreme value type 1 distribution. This assumption is key since it allows market shares to have a closed-form solution.

$$A_{jm}(x, p_{.m}, \delta_{.m}; \theta_2) = \{D_i, v_i, \epsilon_{im} | u_{ijm} > u_{ihm} \forall h = 0, 1, ..., J\} \tag{7}$$

where $p_{.m} = (p_{1m}, ..., p_{Jm})'$ and $\delta_{.m} = (\delta_{1m}, ..., \delta_{Jm})'$. The market share of product j can be written as a function of average utility levels:

$$s_{jm}(x, p_{.m}, \delta_{.m}; \theta_2) = \int_{A_{jm}} dP^*(D, v, \epsilon) = \int_{A_{jm}} dP^*(D) dP^*(v) P^*(\epsilon) \tag{8}$$

With the aforementioned assumption about ϵ, we have that $s_{ijm} = exp(\delta_{jm} + \mu_{ijm})/(1 + \sum_{k=1}^{J} exp(\delta_{km} + \mu_{ikm}))$ is the probability of individual i of buying variety j. In this context, each individual may have a different price sensitivity, and substitution patterns between brands are not derived from functional form. The estimation strategy is to select parameters that minimize the distance between the predicted market share in equation 8 and the observed market share. Equation 8 does not have a closed analytical form, so the integral must be computed using simulation methods.

4.2 Supply

4.2.1 Pricing equations

Following Kim and Cotterill (2008), we assume that there are F firms and each one produces a subset of the varieties $1, \ldots, J$. We assume that the firms solve for each market m an independent maximization problem and that the marginal costs mc vary between markets. Each firm f maximizes profits in market m:

$$\pi_f^m = \sum_{j=1}^{J_f} (p_{jm} - mc_{jm}) \times M \times s_{jm}(\mathbf{p}) - C_f \tag{9}$$

being mc_{jm} the marginal cost of variety j in market m, M the size of the market, $s_{jm}(\mathbf{p})$ the market share of variety j in market m (which depends on the price of all varieties) and C_f the fixed cost of production.

Assuming positive prices and the existence of a pure-strategy Nash-Bertrand equilibrium in prices (Nevo (2001)), the first-order conditions with respect to the prices of the problem are the following set of J equations (we omit the reference to the market m):

$$\sum_{k=1}^{J_f} (p_k - mc_k) \partial s_k(\mathbf{p}) / \partial p_j + s_j(\mathbf{p}) = 0 \tag{10}$$

In vector notation, the first-order conditions became:

$$(\mathbf{p} - \mathbf{mc}) \Delta(\mathbf{p}) + \mathbf{s}(\mathbf{p}) = 0 \tag{11}$$

where $\mathbf{p}$ is the vector of prices for all varieties, $\mathbf{mc}$ is the vector of marginal costs for all varieties, and $\mathbf{s}(\mathbf{p})$ is the vector of market shares. Finally, Δ is a J*J matrix defined in a different way depending on the type of competition that we suppose to exist in the market. If we assume that there is Nash Bertrand competition, that is, that the firms choose their prices simultaneously and in an uncoordinated manner, the matrix is $\frac{\partial s_k(p)}{\partial p_j}$ valid when varieties k and j are produced by the same firm and 0 in the rest of the cases. In other words, the firm behaves like a monopolist with respect to its varieties. On the other hand, if collusion is assumed to exist, the matrix is $\frac{\partial s_k(p)}{\partial p_j}$ valid for the varieties of colluding firms and 0 otherwise. In our case, we build a scenario with perfect price collusion (or monopoly), where the final structure is the joint profit-maximization of all the brands.

Returning to formula 11, we can rewrite solving for the marginal cost for each variety in each market:

$$\hat{\mathbf{mc}} = \mathbf{p} + \Delta(\mathbf{p})^{-1} \mathbf{s}(\mathbf{p}) \tag{12}$$

It is observed that the estimated marginal cost depends on prices and market shares, which are observed, and on the parameters of the demand system $\Delta(\mathbf{p})$, which are estimated. Therefore, we can obtain the marginal costs (and profit margins) from the demand information, making assumptions about how the varieties of each product compete in the market, without looking at the cost information.

4.2.2 Pass-through equations

For the pass-through ratios, following Kim and Cotterill (2008), if there is an industry-wide shock to the marginal cost, the market prices will converge to a new equilibrium, which can be solved as the system of first-order condition in the Equation 11. Now we know the marginal cost (given the size of the shock and the assumption about the model of competence in the market), and the primitives of demand, and we can recover the equilibrium prices. The price pass-through rate is defined as the ratio of the price change to the change in marginal cost:

$$Pass\ through\ rate = (\Delta p / \Delta mc) \times 100 \tag{13}$$

where Δp is the difference between the new equilibrium price that solves the system 11 for the new marginal cost and the old price and $\Delta mc = mc_{new} - mc_{old}$.

4.3 Identification

These estimations must deal with the challenge of controlling for the correlation between prices and the error term, which includes unobservable product characteristics that are observed by consumers but not by the econometrician. As stated in Kim and Cotterill (2008), it is reasonable to think that this correlation is positive because higher levels of unobservable quality of the products can generate that consumers are willing to pay higher prices, and suppliers can set higher prices.

To control for the endogeneity of prices, one needs to find variables that are correlated with prices but are independent of unobserved product characteristics. A set of instruments with a range at least equal to the dimension of the vector of parameters to be estimated is required. In some cases, such as Berry (1994) and Berry et al. (1995) the endogeneity problem is addressed by assuming that the location of brands in the space of product characteristics is exogenous, or at least predetermined. In this context, the characteristics of the other products can be a valid instrument. In this study, as in that of Nevo (2001), there is no variability in the observed characteristics of each brand over time or between locations, so this type of instrument should be ruled out.

In this context, the set of instruments proposed in Hausman (1996) is used, that is, the prices in other localities in the region, exploiting the panel structure of the base. For this, the country is divided into 5 regions: Metropolitan: Montevideo, San José, Canelones; East: Maldonado, Rocha, Treinta y Tres, Lavalleja; South-Central: Flores, Florida, Peach; Coast: Colonia, Soriano, Río Negro; North: Artigas, Salto, Paysandú, Rivera, Tacuarembó and Cerro Largo. The average prices of the product in the region for each month are used as instruments, without taking into account the price of the instrumented locality.

The assumption of identification is that by controlling for variety and demographic characteristics, the changes in the valuation of the varieties are independent between localities. Given this assumption, and because the marginal costs of the same variety in different stores are correlated with each other, the prices of the varieties in other localities are valid instruments. The assumption can be violated if there are national or regional demand shocks that modify the unobserved valuation of all varieties in all stores. A type of shock such as the ones mentioned can occur if the producing companies have advertising campaigns that are coordinated between localities and have an effect on the demand for the varieties. Following Nevo (2001), if it is about mature markets with well-established varieties, it is unlikely that there will be shocks of this type, and furthermore, these shocks can be captured by incorporating dummy variables by period, as we do.

The specification includes dummy variables by product variety, because, as indicated in Nevo (2000), in a context in which the observed characteristics of the varieties do not allow adequate capturing of the factors that determine the utility of individuals, the inclusion of variety fixed effects improves the fit of the model. Another reason was stated when describing the instruments used since the inclusion of dummy variables per variety allows capturing the characteristics that do not vary by market and the variety-specific mean of the unobserved components. By including fixed effects per variety, the coefficients associated with the preferences of individuals for the observed characteristics of the varieties cannot be directly identified. Following Nevo (2000) we recover these parameters using a minimum distance procedure developed in Chamberlain (1982).

4.4 Demand estimation

The estimation method is the one proposed by Berry et al. (1995), but with the differences indicated in Nevo (2001). The first of the differences is about the instrumental variables used, which was described in the previous section. In this context, the identification of demand does not require specifying a functional form on the supply side. The other difference was also mentioned in the previous section and refers to the fact that due to the panel structure of the data available, it is possible to control for the unobservable characteristics of the products using fixed effects by brand.

The model is estimated by the generalized method of moments (GMM) exploiting a population moment condition composed of the product of the instrumental variables and the error term. Let be $Z = [z_1, ..., z_M]$ a set of instruments such that $E[Z'.\omega(\theta^*)] = 0$, while θ (a function of the model parameters), is the error term and θ^* is the true value of the parameters. The estimator is $\hat{\theta} = arg\, min_\theta\, \omega(\theta)'ZA^{-1}Z'\omega(\theta)$ where A is a consistent estimate of $E[Z'\omega\omega'Z]$.

Following Berry (1994) and returning to equation 3, the error term can be decomposed into $\xi_j + \Delta\xi_{jm}$, that is, the average valuation at the national level for the unobserved characteristics of the products and a specific deviation of each market with respect to the average. Since fixed effects are included by brand, the error term is defined as $\Delta\xi_{jm}$.

The unobserved characteristics are computed as a function of the data and parameters, starting from the average utility δ_m, solving the following system of equations for each market:

$$s_{.m}(x, p_{.m}, \delta_{.m}; \theta_2) = S_{.m}\ , \quad m = 1, ..., M \tag{14}$$

being $s_{.m}$ the market share defined in equation 8 and $S_{.m}$ the observed market share. In logit with random coefficients, two steps are required to solve this system of equations. In the first place, the left side of the equality is defined according to equation 8, and the integrals that define the market shares are computed by simulation.

Second, using these market shares, the system of equations defined in 14 is inverted. In the case of the model with random coefficients, the system of equations is non-linear, and therefore the inversion of the model must be done numerically. The system of equations can be solved using the contraction mapping proposed in Berry (1994), which is analogous to iterating over the system:

$$\delta_m^{h+1} = \delta_m^h + \log(S_m) - \log(s_m(x_m, p_m, \delta_m^h, P_{ns}; \delta_m))\ , with\ \ m = 1, ..., M\ and\ h = 0, ..., H \tag{15}$$

Where $s(.)$ are the market shares estimated in the first step, H is the smallest integer such that $\|\delta_m^{H+1} - \delta_m^H\|$ is less than a certain tolerance level and δ_m^H is an approximation to δ_m. Once the inversion has been made, the error term is defined as:

$$\omega_{jm} = \delta_{jm}(x, p_m, S_m; \theta_2) - (x_j\beta + \alpha p_{jm}) \tag{16}$$

The estimation is carried out with the "BLPestimatoR" package of R. The details of the estimation are reported in appendix A.2.

5 Results

5.1 Logit model

As a first step, a logit model (without random coefficients) is estimated, which, despite its limitations regarding the substitution patterns it yields, is an adequate point of reference, mainly to analyze the importance of instrumenting the price and the effects of the instruments.

Table 5 presents the results of different specifications of the logit model. The dependent variable in this model is the logarithm of the market share of variety j in month t, minus the logarithm of the market share of the external option for the same month ($\log(S_{jt}) - \log(S_{0t})$). In the 3 specifications, fixed effects by quarter (we include 7 dummies of quarters) are included. In the second and third specifications, it is controlled by fixed effects by variety, while in the first it is controlled by observable characteristics of the products. Finally, all specifications include the logarithm of the average annual sales of the sum of the three products in the stores by locality, as a proxy of store size.

Table 5: Demand estimation. Logit model

VARIABLES	(1) OLS 1	(2) OLS 2	(3) IV	(4) OLS 1	(5) OLS 2	(6) IV	(7) OLS 1	(8) OLS 2	(9) IV
Price	-0.151***	-0.151***	-0.257***	-0.036***	-0.220***	-0.295***	-0.005	-0.375***	-0.424***
	(0.008)	(0.009)	(0.036)	(0.007)	(0.013)	(0.016)	(0.008)	(0.021)	(0.027)
Log of median age	-3.460***	-4.015***	-3.864***	-11.060***	-10.133***	-9.453***	-9.702***	-9.335***	-9.198***
	(0.980)	(0.845)	(0.868)	(0.954)	(0.859)	(0.884)	(0.938)	(0.803)	(0.751)
Log of median education	-1.058***	-0.634**	-0.858***	-0.513*	-0.790***	-0.981***	-2.265***	-2.042***	-2.033***
	(0.314)	(0.264)	(0.331)	(0.276)	(0.239)	(0.283)	(0.262)	(0.236)	(0.242)
Log of av sales	0.432***	0.469***	0.244***	0.639***	0.529***	0.472***	1.339***	1.049***	1.008***
	(0.055)	(0.051)	(0.091)	(0.051)	(0.050)	(0.050)	(0.046)	(0.045)	(0.045)
Log of av home size	-2.387***	-1.723***	-1.909***	1.345***	-0.009	-0.631	-0.082	-0.443	-0.544
	(0.537)	(0.464)	(0.509)	(0.514)	(0.463)	(0.496)	(0.490)	(0.443)	(0.409)
Observations	2,751	2,751	2,751	3,360	3,360	3,360	3,525	3,525	3,525
R-squared	0.254	0.458	0.415	0.174	0.321	0.312	0.332	0.516	0.515
Product characteristics	Yes	No	No	Yes	No	No	Yes	No	No
Brand dummy	No	Yes	Yes	No	Yes	Yes	No	Yes	Yes
Product	Oil	Oil	Oil	Sauce	Sauce	Sauce	Rice	Rice	Rice
1st stage F-test			313			988			1225

Notes. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Dependent variable is $In(S_jt) - In(S_0t))$. All specifications include fixed effects by quarter (7 dummies of quarters). For each product, the first specification includes product characteristics (dummies for soybean, canola and high oleic in oil, type 1 for rice and concentrated for tomato sauce). In the third specification of each product, the instruments are prices in other localities, similar to Nevo (2001). All specification includes as controls demographic characteristics (the logarithm of the median age and years of education in the locality, and the logarithm of the average home size), and the logarithm of the average annual sales of the sum of the three products in the stores by locality, as a proxy of stores size.

Regressions (i) and (ii) are estimated by ordinary least squares. Regression (i) does not include fixed effects by brand, and this results in the error term containing the unobserved characteristics of the products.

Regression (iii) is estimated by least squares in two stages, instrumenting by the average price in the rest of the localities in the region in each month. It is observed that when using instrumental variables, price sensitivity increases for the three products studied. The fact that when controlling for endogeneity, the sensitivity of demand to prices increases is in line with what is theoretically expected, since it is reasonable to expect a positive correlation between prices and unobservable product quality.

In the 3 specifications, demographic characteristics are included as regressors: the logarithm of the median age in the locality, the logarithm of the median years of education by locality, and the logarithm of the average home size. It is understood that the evaluations that consumers make of the different varieties may have a local component that is captured in part by the inclusion of these variables. As pointed out in Nevo (2001), the coefficients associated with the demographic variables in a model of this type capture the change in the valuation of the product relative to the external option as a function of the demographic characteristics. The results suggest that the valuation for the three articles is lower in localities with younger inhabitants and lower educational levels. In the logit model of random coefficients, demographic characteristics are introduced in a more sophisticated way, but this preliminary analysis suggests that it is relevant to take these variables into account.

Observing the F statistics of the first stage of the specifications with instrumental variables, it can be seen that the proposed instruments are jointly significant, and it cannot be rejected that the instrumental variables have joint power. First-stage R-squares are also high, suggesting some statistical power of the proposed instruments. The complete regressions of the first stage for prices as instruments are presented in Appendix A.1. The central elements to retain from the presented logit model are the importance of controlling for the endogeneity of prices and of using demographic variables.

5.2 Random coefficients logit model

Table 6 presents the results of the random coefficient logit model that was described in section 4.1. The predicted market shares are calculated using equation 8. The demographic information used for the random extractions comes from the 2011 census, using 200 extractions per locality.

The first row contains the mean marginal utility for the price, that is, the linear parameter α. The estimated coefficients for the price are statistically significant in all cases, have the expected sign, and are similar in magnitude to those estimated by the logit with instrumental variables.

The results of the random part show the estimated heterogeneity with respect to the means. It is observed that the estimated parameters of the standard deviations of the observed characteristics are not significant, as well as the majority of the interactions with the demographic characteristics. However, for rice, it is observed that more educated people tend to be less price sensitive and more sensitive to type 1 quality, while for tomato sauce more educated people tend to be less price sensitive and more sensitive to the level of concentration of the product.

Table 6: Results from the full model

	Means	Sd	Education	Age
Oil				
Price	-0.400**	0.002	0.161	-0.065
	0.184	0.178	0.198	0.116
Soy	−2.086	-0.077	2.376	
	2.625	16.158	6.575	
Canola	−4.575	-0.556	-0.988	
	3.843	11.390	5.954	
High oleic	2.715	-0.862	0.485	
	3.843	7.124	2.821	
Tomato sauce				
Price	-0.449***	-0.001	-0.202*	0.072
	0.075	0.166	0.104	0.087
Concentrated	5.445*	-0.177	4.476*	
	1.679	20.390	2.430	
Rice				
Price	-0.426***	0.001	-0.102*	-0.256
	0.106	0.485	0.052	0.246
Type 1	4.528***	0.185	3.263*	1.712
	0.589	3.740	1.738	2.177

Notes. Based on 2751 (oil), 3687 (rice) and 3389 (sauce) observations. All regressions include brand and time (7 dummies of quarters) fixed effects and the logarithm of the average annual sales of the sum of the three products in the stores by locality, as a proxy of store size. Estimated by GMM. *** p<0.01, ** p<0.05, * p<0.1. Prices are instrumented with regional price averages for the variety, as described in section 4.3. Linear coefficients of characteristics different from price are estimated from a minimum distance procedure. Type 1 takes the value 1 if the rice variety has more than 90% of entire grains. Parboiled takes the value 1 if the rice variety is partially boiled. Soy and Canola refer to the plant from which the oil is extracted (the third type in the database is sunflower). Finally, high oleic oil is one that contains at least 75% oleic acid in its composition.

The results for the elasticities are presented below. The price elasticities of market shares in a random coefficient model are defined as follows:

$$\eta_{jkm} = \frac{\partial s_{jm} p_{km}}{\partial p_{km} s_{jm}} = \begin{cases} -\frac{p_{jm}}{s_{jm}} \int \alpha_i s_{ijm}(1 - s_{ijm}) d\hat{P}_D^*(D) d\hat{P}_v^*(v) & \text{if } j = k \\ \frac{p_{km}}{s_{jm}} \int \alpha_i s_{ijm}(1 - s_{ikm}) d\hat{P}_D^*(D) d\hat{P}_v^*(v) & \text{otherwise} \end{cases} \tag{17}$$

Substitution patterns are not derived from functional form (as in a logit model), but from differences in price sensitivity between consumers, allowing for flexible substitution patterns.

Tables 7, 8 and 9 present the own and crossed elasticities for the median of all markets. The elasticity of the variety in the row with respect to a change in the price of the variety in the column is presented. High sensitivities of market shares to changes in prices are observed for the three products: for oil, the own elasticities vary from -6.54 to -15.25, for tomato sauce from -3.67 to -17.64 and for rice from -7.56 to -10.75.

To understand the richness of the substitution patterns that the random coefficient logit model yields, it is useful to remember what the elasticities are like in the logit model without random coefficients:

$$\eta_{jkm} = \frac{\partial s_{jm} p_{km}}{\partial p_{km} s_{jm}} = \begin{cases} -\alpha p_{jm}(1 - s_{jm}) & \text{if } j = k \\ \alpha p_{jm} s_{jm} & \text{otherwise} \end{cases} \tag{18}$$

That is, the cross elasticities within the same "column" are all the same. The presented tables illustrate the changes observed in the logit model of random coefficients.

Table 7: Median own and cross price elasticities, oil

	Condesa soja	Diez soja	Demas soja	Optimo canola	Optimo girasol	Optimo girasol altoleico	Revelacion soja	Rio de la plata soja	Uruguay girasol
1 Condesa soja	-9.07	0.11	0.11	0.19	0.38	0.03	0.07	0.13	0.75
2 Diez soja	1.68	-11.06	0.05	0.18	0.39	0.03	0.04	0.17	0.67
3 Demas soja	1.68	0.08	-10.32	0.08	0.37	0.04	0.05	0.11	0.57
1 Optimo canola	1.86	0.08	0.07	-16.86	0.52	0.02	0.06	0.10	2.03
1 Optimo girasol	1.02	0.05	0.07	0.08	-12.50	0.08	0.03	0.08	0.58
1 Optimo girasol altoleico	0.36	0.00	0.03	0.01	0.29	-6.54	0.01	0.03	0.15
4 Revelacion soja	2.03	0.08	0.11	0.16	0.38	0.04	-11.02	0.13	0.74
5 Rio de la plata soja	1.65	0.09	0.09	0.13	0.40	0.04	0.05	-10.45	0.61
1 Uruguay girasol	1.86	0.10	0.09	0.49	0.57	0.03	0.06	0.12	-15.25

Notes. Cell entries i, j, where i indexes row and j column, give the percent change in market share of brand i with a 1% change in price of the good j. Each entry represents the median elasticity for all markets, weighted by the population in each market.

To complete the analysis of the demand results, the diversion ratios for each product are reported, defined as the fraction of consumers who leave product j after a price increase and switch to product k. As pointed out in Conlon and Mortimer (2021), while own-price elasticities are informative about the market power of the firm, cross-price elasticities alone are insufficient to understand how close

Table 8: MEDIAN OWN AND CROSS PRICE ELASTICITIES, SAUCE

	Cololo	Conaprole	De Ley	Don Perita	Gourmet	Gourmet Napolitana	Qualitas	Rigby	Rigby Italiana	Big
1 Cololo	-4.72	0.21	0.30	0.02	0.05	0.00	0.05	0.07	0.03	0.10
2 Conaprole	0.04	-17.07	0.25	0.11	2.16	0.02	0.15	0.15	0.05	0.19
3 De Ley	0.05	0.31	-5.98	0.06	0.10	0.03	0.10	0.14	0.07	0.18
4 Don Perita	0.07	1.39	0.56	-9.68	1.11	0.03	0.29	0.31	0.09	0.37
3 Gourmet	0.01	3.90	0.11	0.13	-17.64	0.00	0.16	0.16	0.03	0.17
3 Gourmet	0.00	0.13	0.16	0.02	0.03	-3.67	0.04	0.07	0.03	0.08
3 Qualitas	0.06	1.00	0.67	0.21	0.97	0.05	-9.64	0.29	0.07	0.41
5 Rigby	0.04	0.85	0.74	0.12	1.21	0.07	0.29	-9.42	0.13	0.41
5 Rigby	0.04	0.41	0.86	0.08	0.37	0.09	0.18	0.22	-8.33	0.31
6 Big	0.07	0.72	0.63	0.11	0.48	0.05	0.20	0.23	0.08	-8.94

Notes. Cell entries i, j, where i indexes row and j column, give the percent change in market share of brand i with a 1% change in price of the good j. Each entry represents the median elasticity for all markets, weighted by the population in each market.

Table 9: MEDIAN OWN AND CROSS PRICE ELASTICITIES, RICE

	Aruba patna	Blue patna	Blue patna parboiled	Casarone	Green chef	Saman blanco	Saman parboiled	Saman patna	San jose	Shiva patna
2 Aruba patna	-7.74	0.60	0.08	0.05	2.52	1.27	0.35	0.51	0.18	0.61
1 Blue patna	0.28	-10.04	0.15	0.04	2.56	1.11	0.36	0.38	0.11	0.49
1 Blue patna parboiled	0.20	0.61	-9.98	0.03	2.08	0.83	0.33	0.29	0.08	0.37
3 Casarone	0.42	0.62	0.10	-8.09	2.74	1.29	0.35	0.46	0.13	0.57
1 Green chef	0.32	0.69	0.15	0.05	-8.48	1.15	0.36	0.39	0.12	0.51
2 Saman blanco	0.28	0.64	0.13	0.05	2.60	-9.84	0.36	0.39	0.13	0.45
2 Saman parboiled	0.21	0.52	0.11	0.03	1.84	0.90	-9.07	0.26	0.09	0.26
2 Saman patna	0.37	0.66	0.09	0.05	2.64	1.36	0.40	-10.75	0.20	0.52
4 San jose	0.36	0.59	0.10	0.04	2.76	1.21	0.33	0.40	-8.02	0.69
1 Shiva patna	0.39	0.68	0.13	0.06	2.89	1.14	0.30	0.44	0.15	-7.56

Notes. Cell entries i, j, where i indexes row and j column, give the percent change in market share of brand i with a 1% change in price of the good j. Each entry represents the median elasticity for all markets, weighted by the population in each market.

substitutes two products are, and diversion rates are more appropriate to understand this.

It is observed that the highest diversion ratios occur between products with similar observable characteristics. In the last row of each Table, the diversion ratio of the market share of the external option with respect to a change in the price of the good in the column is observed. For the three products, it can be seen that when faced with increases in the prices of each variety, a relevant fraction of consumers stop buying the good, with the outside option generally being above 40 percent.

Table 10: Diversion ratios oil

	Condesa soja	Diez soja	Demas soja	Optimo canola	Optimo girasol	Optimo girasol altoleico	Revelacion soja	Rio de la plata soja	Uruguay girasol
1 Condesa soja	1.00	0.12	0.11	0.12	0.09	0.08	0.12	0.12	0.12
2 Diez soja	0.01	1.00	0.01	0.01	0.01	0.01	0.01	0.01	0.01
3 Demas soja	0.00	0.00	1.00	0.00	0.00	0.00	0.00	0.00	0.00
1 Optimo canola	0.03	0.02	0.01	1.00	0.01	0.00	0.02	0.01	0.06
1 Optimo girasol	0.03	0.03	0.03	0.03	1.00	0.07	0.03	0.03	0.04
1 Optimo girasol altoleico	0.00	0.00	0.00	0.00	0.00	1.00	0.00	0.00	0.00
4 Revelacion soja	0.00	0.00	0.00	0.00	0.00	0.00	1.00	0.00	0.00
5 Rio de la plata soja	0.02	0.02	0.02	0.02	0.02	0.02	0.02	1.00	0.02
1 Uruguay girasol	0.06	0.04	0.04	0.15	0.05	0.03	0.05	0.04	1.00
Outside good	0.84	0.76	0.77	0.67	0.81	0.79	0.75	0.77	0.75

Notes. Cell entries i, j, where i indexes row and j column, give the fraction of unit sales lost by the product j due to an increase in its price of 1% that would be diverted to the i product. Each entry represents the median elasticity for the markets with all varieties of the product, weighted by the population in each market.

Table 11: Diversion ratios sauce

	Cololo	Conaprole	De Ley	Don Perita	Gourmet	Gourmet Napolitana	Qualitas	Rigby	Rigby Italiana	Big
1 Cololo	1.00	0.00	0.01	0.00	0.00	0.02	0.00	0.00	0.01	0.01
2 Conaprole	0.04	1.00	0.03	0.03	0.09	0.02	0.03	0.03	0.02	0.03
3 De Ley	0.24	0.06	1.00	0.10	0.05	0.19	0.10	0.10	0.15	0.12
4 Don Perita	0.01	0.01	0.02	1.00	0.02	0.01	0.02	0.02	0.01	0.02
3 Gourmet	0.03	0.20	0.05	0.12	1.00	0.03	0.10	0.12	0.06	0.08
3 Gourmet	0.10	0.00	0.02	0.01	0.00	1.00	0.01	0.01	0.01	0.01
3 Qualitas	0.04	0.02	0.04	0.04	0.03	0.03	1.00	0.04	0.04	0.04
5 Rigby	0.05	0.03	0.05	0.05	0.06	0.04	0.05	1.00	0.05	0.05
5 Rigby	0.04	0.02	0.05	0.03	0.03	0.04	0.03	0.03	1.00	0.03
6 Big	0.05	0.03	0.06	0.05	0.04	0.05	0.05	0.05	0.05	1.00
Outside good	0.40	0.61	0.68	0.57	0.67	0.57	0.59	0.59	0.60	0.61

Notes. Cell entries i, j, where i indexes row and j column, give the fraction of unit sales lost by the product j due to an increase in its price of 1% that would be diverted to the i product. Each entry represents the median elasticity for the markets with all varieties of the product, weighted by the population in each market.

5.3 Price-cost margins

Table 13 shows the recovered marginal costs and markups for each product, calculated as the mean for all markets[9] The calculations are made under three different conduct assumptions: Nash Bertrand competition with the current ownership, Nash Bertrand competition with single product ownership (in which the price of each brand is set by a profit-maximizing agent that considers only the profits from that brand), and the collusion assumption. To recover the marginal cost, equation 12 and the

[9]In Tables A.3, A.4 and A.5 of the appendix we present the information by brand.

Table 12: Diversion ratios rice

	Aruba patna	Blue patna	Blue patna parboiled	Casarone	Green chef	Saman blanco	Saman parboiled	Saman patna	San jose	Shiva patna
2 Aruba patna	1.00	0.06	0.05	0.06	0.07	0.07	0.05	0.06	0.06	0.07
1 Blue patna	0.04	1.00	0.04	0.04	0.05	0.05	0.04	0.05	0.04	0.04
1 Blue patna parboiled	0.00	0.00	1.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
3 Casarone	0.00	0.00	0.00	1.00	0.00	0.00	0.00	0.00	0.00	0.00
1 Green chef	0.21	0.22	0.20	0.21	1.00	0.27	0.21	0.24	0.19	0.23
2 Saman blanco	0.14	0.16	0.14	0.13	0.17	1.00	0.15	0.16	0.13	0.13
2 Saman parboiled	0.03	0.04	0.04	0.02	0.04	0.05	1.00	0.03	0.03	0.02
2 Saman patna	0.06	0.06	0.05	0.05	0.08	0.07	0.05	1.00	0.05	0.06
4 San jose	0.01	0.01	0.00	0.01	0.01	0.01	0.00	0.01	1.00	0.01
1 Shiva patna	0.06	0.05	0.04	0.05	0.06	0.05	0.04	0.06	0.05	1.00
Outside good	0.46	0.40	0.43	0.43	0.51	0.42	0.45	0.39	0.45	0.44

Notes. Cell entries i, j, where i indexes row and j column, give the fraction of unit sales lost by the product j due to an increase in its price of 1% that would be diverted to the i product. Each entry represents the median elasticity for the markets with all varieties of the product, weighted by the population in each market.

information that arises from the estimation of demand are used. The markup is calculated as (p-cmg)*100/p.

As expected, the recovered marginal costs under the assumption of Nash Bertrand competition with single-product firms are higher than under Nash Bertrand with the current ownership, and these are higher than under collusion. As a counterpart, the markups are smaller.

The inclusion of an assumption of conduct as Nash Bertrand single-product firms is not based on the fact that it is a case of relevant conduct in itself to be tested because it is known which firms sell each variety of product. Its interest lies in the fact that it makes it possible to distinguish between the market power that firms obtain due to their ability to differentiate products from those of their competitors, with respect to that obtained by owning two products perceived as imperfect substitutes by consumers and charging higher prices to those who would charge two firms that sell the good separately (Nevo (2001) and Slade (2004)).

Therefore, with this information, we can say that if we assume that firms are competing in prices a la Nash Bertrand, on average their margins are 27.9% for oil, 36.1% for rice, and 22.3% for sauce. But also we can decompose these margins in two sources: for oil, 18.2 of this 27.9 (65% of the market power) is explained by the ability of the firms to offer products perceived as different from the rest of the market by the consumers, while 35% is explained by the fact that firms own several varieties of the product. If we apply the same reasoning for rice and sauce, 49% of the market power in the rice market is explained by product differentiation and 51% for ownership structure, and for sauce 65% for differentiation and 35% for ownership.

Without additional information, in principle, neither Nash Bertrand competition with current ownership nor collusion can be ruled out. In order to deepen the analysis and rule out some of the behavior assumptions, it is necessary to use information on costs or markups from another source (as in Nevo (2001) or Slade (2004)). If we rely on the observed information on costs and margins on the production side (Table 1), we could say that for the three products, the observed margins are between the estimates assuming collusion and the estimates assuming Nash Bertrand with the observed ownership structure, but closer to the competition à la Nash Bertrand. The literature has broadly followed two paths to deepen the analysis: 1- test each assumption of behavior against the observed data (menu approach, Nevo (2001), Berto Villas-Boas (2007)) or, 2- based on the observed data, recover behavioral parameters, that is, instead of testing whether they are colluding completely or competing completely a la Nash Bertrand, parameters on the degree of collusion are recovered (conduct approach, Miller and Weinberg (2017), (Miller et al. (2021)). However, in this application, no conduct tests will be carried out, due to the problems presented by the "observed" cost information.

Table 13: Mean prices, marginal cost and margins

	(1) Oil mean	(2) Rice mean	(3) Sauce mean
Prices	55.9	32.4	43.9
Single product marginal cost	52.2	26.9	37.9
Current ownership marginal cost	46.4	21.1	34.7
Collusion marginal cost	35.6	9.9	23.0
Single product margin	18.2	17.6	14.5
Current ownership margin	27.9	36.1	22.3
Collusion margin	57.6	74.8	51.4
Observations	2751	3525	3360

Notes. Presented are means of the brand-locality-quarter observations, weighted by the sales. Margins are defined as (p-mc)/p. Marginal cost and margins computed based on the full model reported on Table 6

Table 14 shows the results of price to cost pass-through rates, for every product under every assumption about the behavior of the producers in the market. Under single-product Nash Bertrand competence, the pass-through rates are 51.6% for oil, 44.3% for sauce and 56.5% for rice. As is expected and discussed in the introduction, under the collusion assumption the pass-through rates are the smaller on average for the three products, being 21.6% for oil, 10.6% for sauce and 21.4% for rice. Finally, also as expected, the pass-through rates with the observed ownership structure are in between those for single product and collusion assumptions. The exercise results indicate lower average pass-through rates than those predicted by linear demand with homogeneous products (of 100%). In addition, in general terms, the pass-through is also low when is compared to Table 8 of Kim and Cotterill (2008)[10]. This confirms the intuition discussed in the introduction, related to the fact that in economies with low intensity of competition, it is expected to find higher price levels and markups, which may cause cost shocks to be transmitted to final prices to a lesser extent.

Table 14: Pass-through rate (%)

	Single product	Current ownership	Collusion
Oil	51.6	21.6	21.6
Sauce	44.3	43.3	10.6
Rice	56.5	52.3	21.4
MC shock	10	10	10

Notes. Presented are means of the brand-locality-quarter observations, weighted by the sales. pass-through rate defined as $\Delta p/\Delta mc$.

6 Conclusions

This paper estimates a demand system for differentiated products for oil, tomato sauce, and rice in small and medium-sized retailers. The estimates are used to compute marginal costs, margins, and pass-through ratios from cost to prices that are feasible under different assumptions about how producers compete in these markets. The work seeks to provide empirical evidence on price formation at the microeconomic level in Uruguay, market power and its origin, as well as the ability of producers to pass cost shocks to the final price of the item.

Regarding the elasticities of demand, consumers are highly sensitive to price increases, and substitution patterns between varieties are intuitive. On the other hand, it is observed that most of the decreases in the market share of a variety due to the rise in prices do not translate into increases in another of the varieties for which there is information, but instead, they stop buying those varieties in retailers for which information is available.

In relation to markups, similar levels are found for oil and tomato sauce, around 25% if competition is assumed to be a la Nash Bertrand with the observed ownership structure and 50% with collusion. For rice, there are higher margins, 36% under Nash Bertrand and 75% under collusion.

[10]This comparison is only as a general reference for another basic good.

The exercise carried out allows us to conclude that of the total margin that producers obtain under the Nash Bertrand competition assumption, approximately 65% is explained by their ability to differentiate products for oil and sauce, while 49% is explained by this reason in the case of rice, while the rest is explained by the ownership structure. As indicated in the analysis carried out on these markets by Czarnievicz and Zipitría (2018), market power can have different origins, and it is key from the regulator's point of view to distinguish them, because some of them, such as those that originate in the capacity of the producer of offering differentiated varieties are neither illegal nor undesirable from the perspective of social welfare. Therefore, the decomposition that is carried out in the present work of the origin of market power is relevant when thinking about policies for the promotion and defense of competition.

Regarding the pass-through from costs to prices, there are higher levels under Nash Bertrand competition than under collusion, but in both cases, they are relatively low levels, which in no case exceed 55%. These results are consistent with the intuition that in more concentrated markets and with high levels of market power, low levels of pass-through can be expected.

Finally, it is necessary to mention as limitations that the methodology and the available data do not allow determining what type of competition occurs in the markets, nor what type of interactions occur between producers, distributors, and retailers. For its part, the question of how competition between retailers occurs and how consumers choose and substitute between retailers is also open, given that the unit of analysis used is the locality. But despite these limitations, it is considered that the present analysis allows a deeper understanding of market power and price formation in retail markets using information that is usually available to the commission for the promotion and defense of competition and a methodology that is standard within the industrial organization literature.

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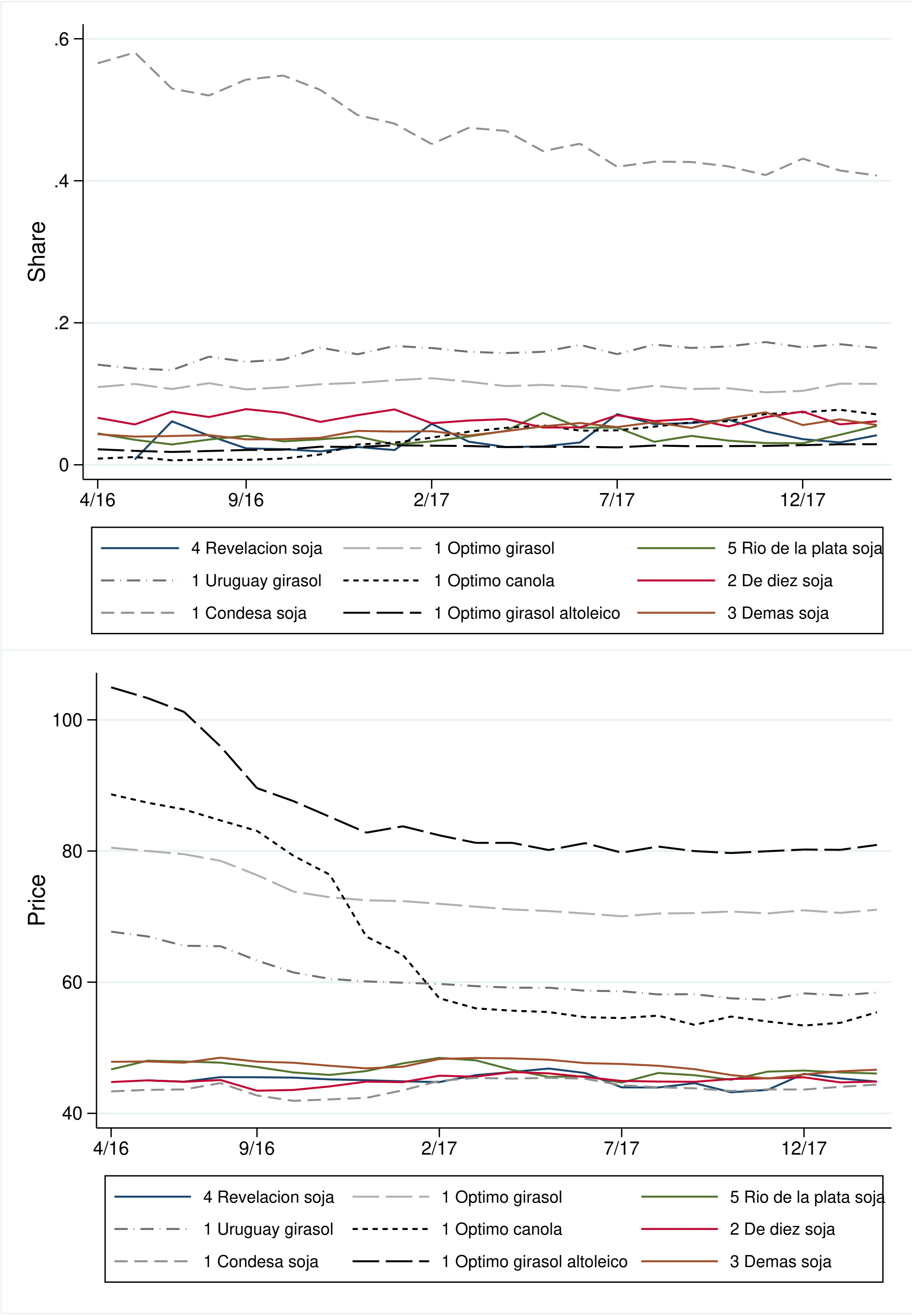
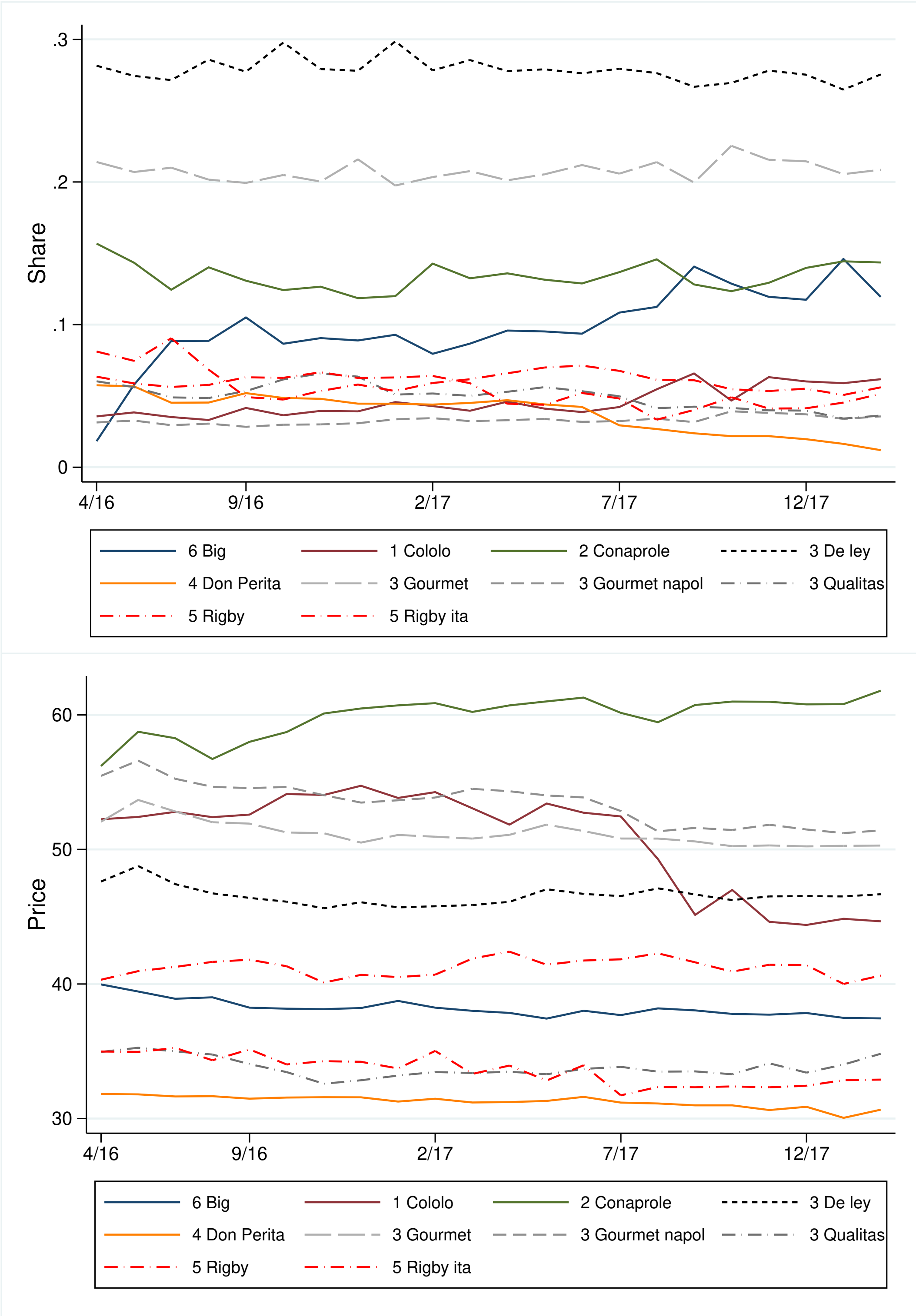
A Appendix

Figure A.1: Evolution of shares and prices by brand: oil



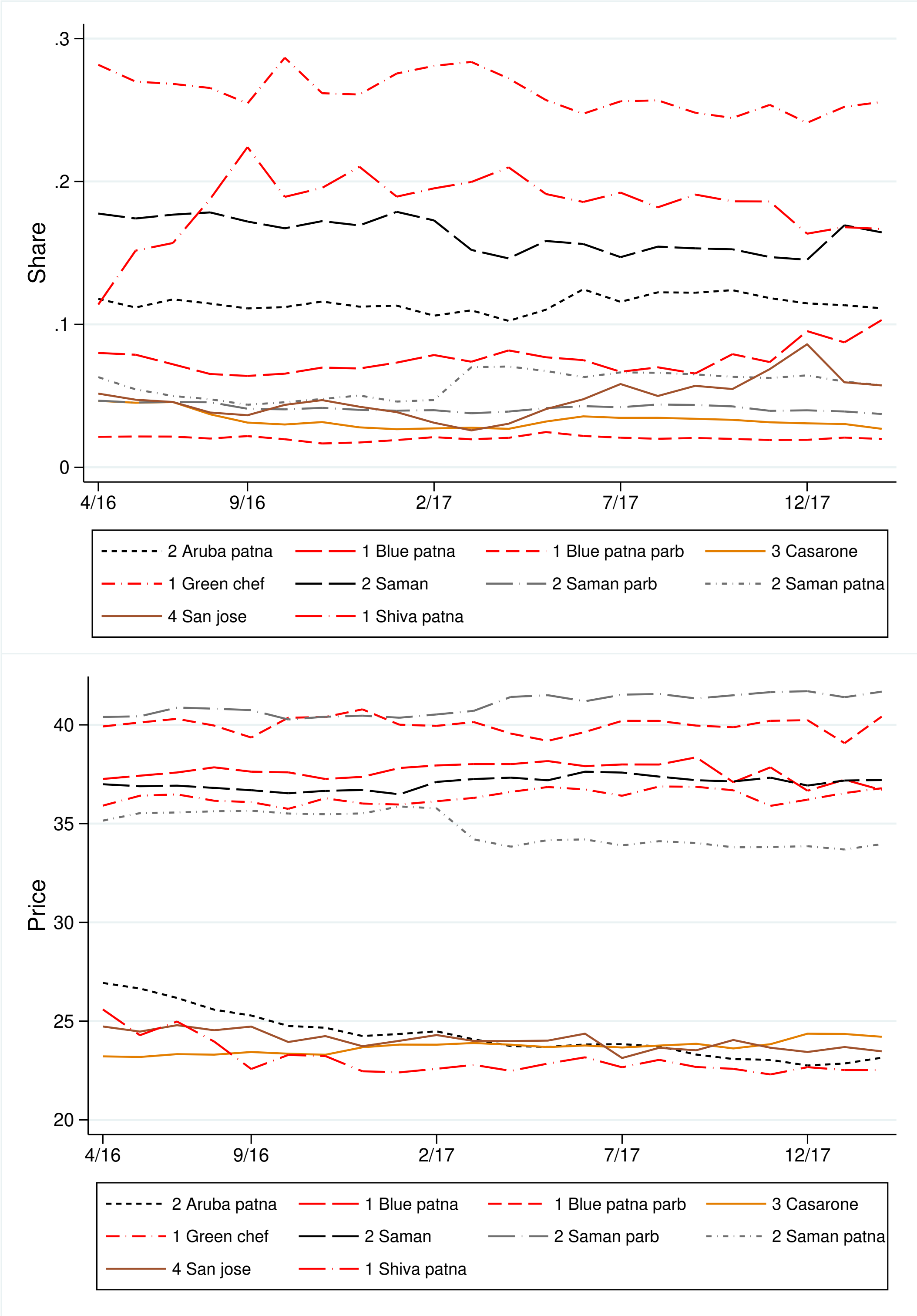
Notes. The upper graph shows the monthly evolution of the shares in all available stores, of each variety of the product with at least a 2% of share in the study period. The lower graph shows the evolution of the average price (per kilo in uruguyan pesos) in all available stores, of each variety of the product in the period studied.

Figure A.2: Evolution of shares and prices by brand: sauce



Notes. The upper graph shows the monthly evolution of the shares in all available stores, of each variety of the product with at least a 2% of share in the study period. The lower graph shows the evolution of the average price (per kilo in uruguyan pesos) in all available stores, of each variety of the product in the period studied.

Figure A.3: Evolution of shares and prices by brand: rice



Notes. The upper graph shows the monthly evolution of the shares in all available stores, of each variety of the product with at least a 2% of share in the study period. The lower graph shows the evolution of the average price (per kilo in uruguyan pesos) in all available stores, of each variety of the product in the period studied.

Table A.1: DESCRIPTIVE STATISTICS OF DEMOGRAPHICS

Departamento	Localidad	Municipio	Age	Education years	Sex
Artigas	Artigas		45.91	8.684	0.455
Artigas	Bella Union		44.76	8.305	0.507
Canelones	Barros Blancos		43.29	7.678	0.478
Canelones	Canelones		46.06	8.961	0.464
Canelones	La Paz		44.92	8.688	0.454
Canelones	Las Piedras		43.92	8.214	0.443
Canelones	Pando		43.90	8.575	0.480
Canelones	Parque Del Plata		47.59	9.508	0.447
Canelones	Paso Carrasco		43.04	8.821	0.487
Canelones	Pinar		43.67	10.46	0.491
Canelones	Progreso		43.56	7.864	0.480
Canelones	Salinas		46.75	9.780	0.468
Canelones	San Ramon		46.75	7.972	0.499
Canelones	Santa Lucia		46.43	8.699	0.474
Canelones	Sauce		45.54	8.363	0.469
Canelones	Solymar		46.04	10.93	0.466
Canelones	Toledo		42.67	8.001	0.471
Cerro Largo	Melo		46.87	8.392	0.442
Colonia	Carmelo		47.23	8.436	0.466
Colonia	Colonia		46.64	9.127	0.454
Durazno	Durazno		46.22	8.534	0.465
Flores	Trinidad		48.69	8.035	0.500
Florida	Florida		47.19	8.720	0.453
Lavalleja	Minas		48.57	8.523	0.450
Maldonado	Maldonado		44.67	8.553	0.468
Maldonado	Piriapolis		48.53	9.397	0.470
Maldonado	Punta Del Este		47.46	11.98	0.465
Maldonado	San Carlos		44.03	8.524	0.457
Montevideo	Montevideo	A	45.83	8.123	0.454
Montevideo	Montevideo	B	43.90	12.21	0.449
Montevideo	Montevideo	C	47.37	11.01	0.451
Montevideo	Montevideo	CH	48.02	13.07	0.437
Montevideo	Montevideo	D	45.32	8.490	0.444
Montevideo	Montevideo	E	47.31	11.62	0.486
Montevideo	Montevideo	F	44.06	8.211	0.471
Montevideo	Montevideo	G	46.38	8.897	0.424
Paysandu	Paysandu		46.62	9.071	0.472
Rio Negro	Fray Bentos		45.36	8.844	0.481
Rio Negro	Young		44.51	7.978	0.529
Rivera	Rivera		45.52	8.432	0.433
Rocha	Rocha		46.88	8.710	0.456
Salto	Salto		44.77	8.739	0.467
San Jose	Ciudad Del Plata		41.76	7.740	0.468
San Jose	Libertad		45.09	7.983	0.470
San Jose	San Jose De Mayo		46.42	8.555	0.484
Soriano	Dolores		45.12	8.309	0.476
Soriano	Mercedes		46.46	8.997	0.457
Tacuarembo	Paso De Los Toros		46.97	7.785	0.468
Tacuarembo	Tacuarembo		45.53	8.298	0.436
Treinta Y Tres	Treinta Y Tres		47.46	8.862	0.474

Notes. Average age, education years and sex by region, obtained from Census 2011.

A.1 First stage of logit model

Table A.2: First stage logit model

VARIABLES	(1) Price	(2) Price	(3) Price
1 instrument	0.445	-1.358**	-0.925
	(1.659)	(0.617)	(0.757)
2 instrument	4.180**	-0.472	0.184
	(1.954)	(0.916)	(0.870)
3 instrument	-2.835	-2.300**	-0.924
	(2.281)	(1.119)	(0.854)
4 instrument	-0.242	-2.603**	-0.131
	(2.109)	(1.053)	(0.833)
5 instrument	5.547***	0.126	-0.982
	(1.748)	(0.988)	(0.799)
6 instrument	-0.337	-4.553***	-1.382*
	(1.753)	(1.020)	(0.740)
7 instrument	1.671	4.101***	-1.207
	(1.655)	(1.354)	(0.783)
8 instrument	-2.049	-2.596**	1.204
	(1.556)	(1.193)	(0.985)
9 instrument	-0.717	-4.024***	0.477
	(1.540)	(1.105)	(1.086)
10 instrument	0.024	-1.671	-1.187
	(1.494)	(1.240)	(1.006)
11 instrument	-4.362*	-0.551	-5.691***
	(2.639)	(1.334)	(1.012)
12 instrument	3.932	-2.726**	1.724*
	(2.394)	(1.285)	(0.984)
13 instrument	4.796*	0.689	-2.821***
	(2.723)	(1.047)	(0.819)
14 instrument	-6.215**	-9.480***	3.870***
	(2.700)	(1.214)	(0.762)
15 instrument	-9.771***	0.672	-1.565*
	(3.138)	(1.244)	(0.893)
16 instrument	6.969*	-3.335***	-1.694*
	(3.675)	(1.039)	(0.969)
17 instrument	-4.161	-0.796	-1.092
	(3.364)	(0.986)	(0.930)
18 instrument	8.269***	-5.024***	-1.944**
	(3.007)	(1.105)	(0.982)
19 instrument	-7.032**	-0.798	3.950***
	(2.856)	(1.169)	(0.977)
20 instrument	3.507	-1.116	-1.770*
	(2.396)	(1.235)	(1.050)
21 instrument	-3.059*	-1.131	-3.549***
	(1.613)	(1.374)	(0.822)
22 instrument	-4.876**	-1.477	-2.410***
	(2.288)	(1.218)	(0.665)
Observations	2,751	3,360	3,525
R-squared	0.826	0.926	0.937
Product	Oil	Sauce	Rice

Notes. First stage of IV regressions reported in Table 5

Table A.3: Margins by product for oil

	Single product	Current ownership	Collusion
Condesa soja	22.2	36.4	55.8
Diez soja	12.8	12.8	23.3
Demas soja	98.1	98.1	192.5
Optimo canola	8.4	32.3	49.7
Optimo girasol	10.6	24.3	36.0
Optimo girasol altoleico	-22.6	-8.8	1.2
Revelacion soja	12.1	12.1	57.2
Rio de la plata soja	11.3	11.3	51.2
Uruguay girasol	11.4	27.5	42.5

Notes. Presented are means of the brand-locality-quarter observations, weighted by the sales. Margins are defined as (p-mc)/p. Margins computed based on the full model reported on Table 6.

Table A.4: MARGINS BY PRODUCT FOR SAUCE

	Single product	Current ownership	Collusion
Cololo	6.0	6.0	37.5
Conaprole	9.2	9.2	31.3
De Ley	24.9	35.8	50.0
Don Perita	12.3	12.3	61.0
Gourmet	16.9	28.4	42.1
Gourmet napolitana	16.9	35.7	50.5
Qualitas	13.7	42.3	62.4
Rigby	16.4	19.8	70.5
Rigby italiana	15.1	19.8	60.0
Pure de tomate Big	13.9	13.9	54.7

Notes. Presented are means of the brand-locality-quarter observations, weighted by the sales. Margins are defined as (p-mc)/p. Margins computed based on the full model reported on Table 6.

Table A.5: Margins by product for rice

	Single product	Current ownership	Collusion
Aruba patna	16.2	35.3	72.8
Blue patna	12.7	30.4	44.5
Blue parna parboiled	11.0	29.6	41.1
Casarone	14.2	14.2	67.7
Green chef	21.2	31.0	45.6
Saman blanco	15.0	21.3	44.8
Saman Parboiled	12.2	21.6	42.5
Saman patna	12.2	24.4	49.4
San jose	15.0	15.0	68.8
Shiva patna	20.3	46.5	68.8

Notes. Presented are means of the brand-locality-quarter observations, weighted by the sales. Margins are defined as (p-mc)/p. Margins computed based on the full model reported on Table 6.

A.2 Appendix: Demand estimation details

The implementation of the estimation of the model proposed by Berry et al. (1995) requires determining a method to approximate the integral, an optimization algorithm, initial values, and convergence criteria. Brunner et al. (2017) discuss implementation alternatives so that the estimation results are adequate and the R package "BLPestimatematoR" is provided, which is used in the present work to carry out the estimation.

Regarding the simulation to approximate the integral of the market shares, 200 MLHS (latin hypercube sampling draws) draws are used. The sensitivity of the results to increasing the number of extractions to 1000 is tested, corroborating that there are no relevant differences in the results. The number of extractions cannot be greater than the number of extractions of the observable characteristics of the individuals. The algorithm used for optimization is BFGS (Broyden–Fletcher–Goldfarb–Shanno).

Regarding the iterations of the contraction, a maximum of 5000 iterations is set or until it is less than 1e-06. Finally, following Nevo (2000) and Chidmi and Lopez (2007), as starting guesses for the average utility vector (γ) the results obtained in the logit model are used.

Property Rights and Effort Supply*

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Abstract

Direct evidence on how effort provision varies across different ownership structures remains scant. In this paper we investigate the absence behaviour of individuals employed in worker cooperatives, that is, in firms owned and ultimately controlled by their workforce. Leveraging monthly employment data matched with certified sick leave records and exogenous variation in the generosity of the Uruguayan paid sick leave regime, we show that both the incidence and duration of sickness-related absences differentially increased for individuals affected by the reform and employed in worker cooperatives. The effect is driven by members' behaviour, both short-term and long-term absences, hard-to-diagnose (and, hence, more prone to moral hazard reporting problems) musculoskeletal conditions, and individuals employed at medium-sized and large cooperatives. We also find suggestive evidence that conventional firms used dismissals more actively than cooperatives as a threat to keep absenteeism in check after the reform. Complementary survey evidence shows that concerns about work ethics became increasingly salient among managers of large cooperatives. Small cooperatives did not experience a similar escalation of absence behaviour. This group of cooperatives seems to rely on a distinct workplace discipline environment based on peer monitoring and less hierarchical supervision.

JEL Classification: I18, J22, J54

Keywords: effort, absenteeism, sick pay, cooperatives, property rights, teams, moral hazard.

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1 Introduction

The impact of property rights systems on workers' behaviour is a topic of perennial interest to economists and organisational scholars (Alchian and Demsetz 1972; Hart and Moore 1998; Hansmann 1996). However, direct evidence concerning individuals' effort responses under different ownership structures remains rare.

This paper fills this gap by assessing differences in absence behaviour, a proxy of effort choices, across individuals employed under two sharply distinct contractual arrangements: worker cooperatives and conventional investor-controlled firms. Worker cooperatives are enterprises in which the workforce has ultimate control rights (Dow 2003). Their members usually own and manage the company on a 'one person, one vote' basis, regardless of the amount of capital they supply to the cooperative. These organisational features stand in sharp contrast to those exhibited by conventional firms, in which outside owners hire labour, appoint managers and have the right to appropriate the residual income. Worker cooperatives are diffused in certain European regions, such as the Basque Country and Emilia Romagna. During specific periods, cooperatives also played a prominent role in certain sectors, such the US plywood industry (Pencavel 2001). Related arrangements, such as partial employee ownership schemes (e.g. ESOPs) and professional partnerships, are also common in the US and Europe.[1]

Measures of work effort are hard to observe, given the team-based nature of most production settings. Absenteeism is an important dimension of workers' behaviour that can be measured at the individual level. It represents a form of employee withdrawal behaviour that can be costly for firms and organisations. Firms may suffer from productivity losses and incur extra costs from employing temporary workers or from paying regular workers overtime in order to cover for absent employees (e.g. Herrmann and Rockoff 2012).[2] Interestingly, arguments concerning work incentives in cooperative firms date back to early economic writings.[3] For example, John Stuart Mill and Alfred Marshall highlighted potential advantages of worker cooperatives:

> "the general sentiment of the community, composed of the comrades under whose eyes each person works, would be sure to be in favour of good and hard working, and unfavourable to laziness, carelessness, and waste." (J. S. Mill, 1879, pp. 518-519).
>
> "[Cooperatives] render unnecessary some of the minor work of superintendence that is required in other establishments; for their own pecuniary interests and the pride they take

1. In the US, Kruse 2022 reports that about 20% of private sector employees own company stock. Roughly 12% of European companies (100+ employees) had put in place an employee ownership plan in 2015, ranging from 7% in The Netherlands to 25% in Belgium (Ligthart, Poutsma, and Brewster 2022)

2. Hensvik and Rosenqvist 2019 show that the extent of production disruptions due to absenteeism depends on firms' ability to find internal substitutes for absent workers.

3. Quotes are taken from Jones 1976.

in the success of their own business make each of them averse to any shirking of work either by himself or by his fellow-workmen." (A. Marshall, 1964, pp. 254-255).

Instead, Sidney and Beatrice Webb, co-founders of London School of Economics and Political Science, raised concerns about the relationship between management and members in this type of firms:

> "The relationship set up between a foreman or manager, who has throughout the working day to give orders to his staff, and the members of that staff who, assembled in general meeting, criticise his action or give him directions, with the power of dismissing him if he fails to conform to their desires, has always been found to be an impossible one" (S. and B. Webb, 1920, p.166).

From the perspective of modern economic analysis, the impact of cooperative property rights on absence behaviour is theoretically ambiguous and remains an open empirical question. On the one hand, several explanations point to weaker work incentives and greater incidence of workers' absenteeism in cooperatives. First, cooperative teams may suffer from the classical free rider problem (Alchian and Demsetz 1972). This may be exacerbated by the *de facto* job security enjoyed by cooperative members, limiting the scope for using the threat of dismissal as a mechanism to keep shirking behaviour in check. Second, managerial discretion to impose sanctions and dissolve labour contracts may be more limited in cooperatives than in conventional firms (Hart and Moore 1998). Indeed, worker cooperatives are characterised by a dual-authority structure. Worker-principals appoint managers, set objectives and monitor the implementation of firm policies. In turn, managers, acting as quasi-principals, organize and monitor the production process and the actions of the workers. Interestingly, while workers have the power to dismiss managers, managers cannot replace workers without consulting the membership (Ben-Ner, Montias, and Neuberger 1993). Finally, egalitarian compensation policies implemented by cooperatives may induce negative selection of workers both at the bottom and the top of the ability distribution, distorting incentives of frontline workers and managerial quality (Kremer 1997; Abramitzky 2009; Burdin 2016).[4]

On the other hand, the fact that cooperatives rely more extensively on group-based profit sharing and on team-based work may mitigate absence behaviour driven by moral hazard. Profit-sharing makes workers residual claimants on the income stream resulting from the noncontractible effort supplied to the firm. This may provide an incentive to reduce absences, particularly in small cooperatives. Moreover, horizontal peer pressure and social emotions may help to save on monitoring inputs, sustain

4. Workers' experience in cooperatives may be more intense and stressful than in a conventional business as members have both production and decision-making responsibilities. This suggests that cooperatives, far from being idyllic workplaces, may be better described as "high-expectation, high-stress work systems" (Arando et al. 2015).

high-effort norms and curb absenteeism in cooperative teams (Putterman 1984;Kandel and Lazear 1992; Hamilton, Nickerson, and Owan 2003; Putterman 1984; Carpenter et al. 2009).[5] As the entire cooperative team suffers when one worker-member is absent from work, the returning team member can be exposed to informal group sanctions.[6] Although profit sharing provides weak incentives to work harder in large organisations, it might suffice to induce reciprocal workers to report each other for shirking (Carpenter, Robbett, and Akbar 2018). Finally, shirking on effort can be deterred in cooperative teams by relying on repeated game mechanisms as long as members expect to interact in the future and are sufficiently patient (Macleod 1984; Putterman and Skillman 1992; Dong and Dow 1993).

To shed light on this debate, our empirical analysis relies monthly employment history administrative records matched with unique individual-level information on certified sick leave over the period 2005-2013. We exploit variation created by a paid sick leave reform that increased the generosity of sickness insurance for certain workers in Uruguay. The reform gradually increased the sick pay cap, providing exogenous variation in sick leave compensation across individuals depending on their pre-reform wage. This setting allows us to use a difference-in-differences (DiD) approach, including heterogeneous treatment effects in order to capture the differential response of workers employed in cooperatives relative to individuals employed in conventional private-sector firms.

The analysis yields two basic results. First, we find that the increase in sick leave pay raised the probability of being absent from work in a given month by 1.6 percentage points more among treated individuals employed in cooperatives than among treated individuals employed in conventional firms. Second, the duration of sickness-related absence spells for treated cooperative members increased by 0.4 days relative to the other groups in a given month. In relation to the pre-reform situation of treated individuals employed in worker cooperatives, sickness absences in the extensive and intensive margins increased by 40% and 55%, respectively. Results from an event-study analysis suggest that the absence behaviour of these individuals was on a similar pre-reform trend relative to the other group. By excluding workers who switched between organisational forms during the period, we show that the results are not driven by non-random sorting into cooperatives due to the reform. The fact that we observe a similar trend in absence behaviour over a period of six years before the reform also suggests

5. Cooperative behaviour in public good games can be sustained by relying on social punishment (Fehr and Gächter 2000). However, peer sanctions may also be targeted at high-contributors (Herrmann, Thöni, and Gächter 2008; Ertan, Page, and Putterman 2009).

6. The cost to the organisation when a worker shirks by being absent and taking excessive paid sick leave may be less salient in the Uruguayan context as the Uruguayan regime has no experience rating sick leave insurance (i.e the payroll tax rate does not rise when more of the firm's workforce receives paid sick leave).

that a more general pattern of selection of absence-prone individuals into cooperative is unlikely to explain our findings.

Using our DiD framework, we explore several potential mechanisms that may account for the observed differences in absenteeism: (1) the differential shift in absence behaviour among treated individuals employed in cooperatives is explained by both short-term and long-term absences, suggesting that this type of firms not only face potential moral hazard problems but also facilitate greater take-up of sick leave motivated by genuine health problems; (2) the increase in absenteeism is entirely driven by cooperative members (no significant effects are obtained when the analysis is restricted to employees in conventional firms and cooperatives); (3) there is no differential increase in extended weekend absences ('Monday effect'); (4) the analysis of disease-specific behavioural responses reveals a differential increase in hard-to-diagnose (and more prone to moral hazard reporting problems) musculoskeletal conditions for treated individuals employed in cooperatives; (5) the dynamics of layoffs suggests that conventional firms use the threat of dismissal more actively than worker cooperatives as to keep absenteeism in check after the reform; and (6) the differential increase in absenteeism is entirely driven by individuals employed in medium-sized and large cooperatives, precisely where one would expect the dilution of work incentives to be more severe.[7].

Complementary survey-based evidence on worker supervision and managers' perceptions, collected before and after the reform, suggests more negative views on absenteeism and work ethics among managers of large cooperatives. Interestingly, small worker cooperatives do not seem to have experienced a similar erosion of work incentives. When the analysis is restricted to the subsample of small firms, our DiD estimates show no differential increase in absenteeism for individuals employed in worker cooperatives after the reform. Moreover, small cooperative exhibit lower supervision intensity than comparable conventional firms and extensively rely on mutual monitoring among coworkers as an alternative discipline device.

The paper contributes to different strands of research. First, we add to a long-standing literature examining moral hazard in team production and how the allocation of controls rights over productive assets affect workers' incentives (Alchian and Demsetz 1972; Holmstrom 1982; Macleod 1984; Macleod 1987; Putterman 1984). Our paper relates to previous research on incentives in communal organisations (Abramitzky 2008, 2009, 2011) and to a series of studies examining the productivity effect of worker cooperatives vis-à-vis conventional firms (Craig and Pencavel 1995; Fakhfakh, Pérotin, and Gago 2012; Pencavel 2013; Monteiro and Straume 2018; Young-Hyman, Magne,

7. Our results are consistent with recent qualitative evidence documenting problems of workplace absenteeism prior to the demise of the world's biggest industrial worker cooperative (Basterretxea, Heras-Saizarbitoria, and Lertxundi 2019)

and Kruse 2022). In a recent paper, Montero 2021 exploits exogenous variation in ownership rights induced by a land reform in El Salvador. Using a regression discontinuity design, he finds that cooperatives are less productive than conventional haciendas when producing cash crops, but more productive when producing staple crops. While previous studies rely on firm-level measures of productivity, our paper is one of the first attempts to provide direct evidence of individuals' effort provision in the form of absenteeism in worker cooperatives.

Second, this paper adds to the literature on sick leave insurance and absence behaviour (Henrekson and Persson 2004; Ziebarth and Karlsson 2010; Ziebarth 2013; Paola, Scoppa, and Pupo 2014; Ziebarth and Karlsson 2014; Pichler and Ziebarth 2017, Bryson and Dale-Olsen 2017; Marie and Vall-Castello 2022). While previous studies have focused exclusively on the U.S. and European countries, little is known about the incentive effects of paid sick leave reforms in less developed countries. Moreover, we contribute to understanding the role of firm organisation in moderating the interplay between sick leave insurance and workplace absenteeism (e.g. Bennedsen, Tsoutsoura, and Wolfenzon 2019). Previous research has analysed the effect of probationary periods (Ichino and Riphahn 2005) and sick leave reforms in the public sector (Paola, Scoppa, and Pupo 2014). According to these studies, workers' behaviour is sensitive to the level of employment protection, sick leave compensation and monitoring intensity. Interestingly, there is extensive evidence documenting greater job security in worker cooperatives compared to conventional firms (Burdín and Dean 2009; Pencavel, Pistaferri, and Schivardi 2006; Garcia-Louzao 2021). The fact that cooperative members "buy" an implicit long-term employment guarantee may have an effect on their absence behaviour. Indeed, our study shows that the impossibility of using dismissal threats as a discipline device seems to be an important channel behind the increase in absenteeism among individuals employed in worker cooperatives.

Finally, our paper contributes to the scant literature on worker voice and non-pecuniary dimensions of jobs, such as health outcomes. Arnold, Brändle, and Goerke 2018 find evidence of higher utilisation of sick leave in German firms with works councils. Exploiting exogenous changes in codetermination rules among Finnish firms, Harju, Jäger, and Schoefer 2021 find no effect of worker voice institutions on sickness-related absences.[8] Goerke and Pannenberg 2015 study the effect of a reduction of statutory paid sick leave using self-reported survey data from Germany. They find a positive relationship between trade union membership and sickness absence and a stronger reaction to the reduction in paid sick leave among union members than among non-members. As the German reform applied across the board to all private workers, their treatment group is entirely composed of private-sector workers and the control group

8. Blasi et al. 2010 find a positive association between employee ownership and absences in the US context.

comprises public-sector workers and self-employed workers. In this paper, we restrict the analysis to private sector workers employed both in worker cooperatives and in conventional enterprises. By relying on high-frequency administrative data, including information on the exact start and end date of each absence spell, our analysis is less affected by the kind of measurement errors that typically pervade survey data. Most importantly, the data allows us to extensively investigate the underlying channels through which the differential response of cooperative members manifests itself. Interestingly, while the studies mentioned above have analyzed institutional arrangements conveying limited power to workers, such as works councils and minority board-level representation, our paper contributes to understanding how the assignment of more extensive control rights to workers affects individual and firm outcomes.

The remainder of the paper is organized as follows. The next section describes the Uruguayan sick leave reform and provides contextual information on worker cooperatives. Section 3 explains the data and the identification strategy. Section 4 presents the main findings, provides evidence concerning identification assumptions and reports results from several robustness checks. Section 5 uncovers different mechanisms that may account for the differential behavioural response of individuals employed in worker cooperatives. Section 6 reports complementary survey evidence on supervision intensity and managerial perceptions about absenteeism and work ethics in cooperatives and conventional firms. Section 7 concludes.

2 Institutional context

2.1 Background on the Uruguayan paid sick leave reform

According to the sick leave legislation in Uruguay, a worker experiencing a sickness episode receives an amount b, which represents a constant replacement ratio (70%) of her last wage (w) up to a maximum benefit amount (bmax), where the replacement rate decreases.[9] The benefit cap is defined in terms of Bases de Prestación y Contribución (BPC), where BPC is the basic unit of measurement used to calculate different social benefits in the Uruguayan social security system.[10] Therefore, the sick leave pay is

9. The fact that the sick leave benefit is a kinked function of previous earnings makes the design of the Uruguayan system comparable to social insurance programs in developed countries, such as the Norwegian public sick leave (Bryson and Dale-Olsen 2019) and unemployment insurance in U.S. states (Landais 2015).

10. 1 BPC is equivalent to 3848 Uruguayan Pesos (USD 117/January 2018). Source: Banco de Prevision Social.

computed according to the following rule:

$$b = \begin{cases} 0.7w & \\ bmax & \text{if } 0.7w > bmax \end{cases} \tag{1}$$

To be eligible, the worker must have worked and paid social security contributions for at least 3 months in the year preceding the illness episode. As is common in other public sick leave regimes, a physician has to certify the worker's health condition. The worker is not entitled to any payment during the first three days of sick leave and can receive the benefit for a maximum of one year; the benefit may be extended for an additional year under special circumstances (Amarante and Dean 2017). The sick leave pay is not disbursed by the employer but by the public health insurance system. The program is funded from general taxation and social security contributions are paid by both employers and employees. In contrast to experience rating insurance systems, employers' payroll tax rates do not depend on the number of workers firms have had on sick leave in the past.

Before the reform, the benefit cap was 3 BPC. Therefore, those workers for whom 0.7w exceeded the threshold of 3 BPC received exactly 3 BPC as paid sick leave. Figure 1 describes the evolution of the paid sick leave schedule over the period analysed in this paper. As a result of the reform, the benefit cap gradually increased by 1 BPC per year starting from January 2011. By January 2013, the last year included in our study, the benefit cap had reached 6 BPC.[11] Figure 2 plots the evolution of the ratio between the benefit cap and the average wage before and after January 2011, confirming the sharp relative increase of the benefit cap. The spikes observed in the data correspond exactly to the reform schedule (January 2011, 2012 and 2013).

2.1.1 Worker cooperatives in Uruguay

Worker cooperatives are defined as enterprises where members jointly carry out the production of goods or services activities and have control over important economic decisions.[12] Usually, members jointly own and manage the firm on a "one person, one vote" basis regardless of their capital contribution and the residual is distributed among them according to a certain sharing rule.

In Uruguay, worker cooperatives are those firms that are legally registered as producer cooperatives (PCs) in which the employee-to-member ratio does not exceed 20%. These firms are allowed to hire salary employees but they must still comply with the legislated maximum percentage of hired workers in order to receive certain tax ad-

11. The reform was fully phased in by January 2015 when the benefit cap reached its current level of 8 BPC.

12. This section draws on Burdin 2016.

vantages – in particular, the exemption from paying the employer payroll tax to social security. The law also requires a minimum of six members to register a new cooperative firm.

Although their key organisational features are predetermined by law, worker cooperatives have discretion over a broad range of associational rules. With respect to governance structure, worke cooperatives must have a general workers' assembly that selects a council to supervise the daily operations (the council, in turn, usually selects the managers). Each member has only one vote, regardless of his capital contribution to the firm. Physical assets can be owned by their members either collectively or individually. Under collective ownership, members do not own tradable shares but enjoy the right to usufruct as long as they work in the firm. Under individual ownership, members own capital shares that vary with the firm's value. Most Uruguayan worker cooperatives operate under a collective ownership regime. As in other countries, membership markets are extremely rare in Uruguay: fewer than 10% of Uruguayan worker cooperatives are owned by their workforce through individual shares (Alves et al. 2012).

3 Data and identification

3.1 Data

Our empirical analysis is based on longitudinal individual-level administrative records from the Uruguayan social security system. The data were provided by Banco de Prevision Social, the agency in charge of social security affairs in Uruguay. Employers are obliged to deliver monthly information on their employees to the agency, which uses that information to calculate pension and social benefits. To conduct this study, we combine three different databases. First, we use monthly employment history data from a random sample of 300,000 individuals who were registered in the social security system for at least one month during the period 2005-2013. The structure of the data is an unbalanced panel of workers, containing information on wages, personal attributes of the worker (gender, age, tenure), and the firm in which she works (firm size, industry, region). Each worker-month observation is associated with a firm identification number so that job changes (or any other discontinuity in the individual's employment history) can be tracked. Moreover, we obtain similar employment history data for the universe of individuals employed in worker cooperatives. Finally, and crucially for the purpose of this study, we match individual-level records of certified sickness absences, including the start and end date of each sickness absence spell, and sick leave payment. Information on short sickness spells (fewer than 4 days) and diagnosis classified according to the International Classification of Diseases (ICD) is only available since 2010. For this reason, our investigation mainly focuses on spells of more than 3 days.[13]

We restrict the sample in several ways. First, we focus on workers employed in non-agricultural private firms, excluding public, rural and construction workers. Second, we only consider eligible individuals, i.e. those who made social security contributions for at least 3 months (or 75 days in the case of day labourers) in the year preceding the sickness spell. The final dataset is an unbalanced panel from January 2008 to December 2013, i.e. three years before and after the sick leave reform. Descriptive statistics for the final sample are presented in Table 1. The resulting sample includes, on average, about 36,965 individuals in each month. The total number of individual-month observations is 2,625,338, corresponding to 52,751 and 3,532 individuals employed in conventional firms and worker cooperatives, respectively. The composition of the two groups is different: individuals employed by worker cooperatives are older than those employed by conventional firms and, in the latter case, the percentage of small firms (less than 20 workers) is higher. Proportionately fewer women are employed by worker cooperatives than by conventional firms, particularly in the treatment group. On average, both the incidence and duration of sickness absences appear to be higher in cooperatives.

3.2 Identification

Before the reform, sick leave pay was subject to a benefit cap equivalent to 3 BPC. In other words, an individual for whom 70% of her total monthly earnings exceeded 3 BPC received exactly 3 BPC. As explained in section 2, the Uruguayan sick leave reform gradually increased this maximum benefit cap starting in January 2011. Our identification strategy exploits the exogenous increase in the generosity of paid sick leave for this group of workers. We compare the evolution of sickness absence (incidence and duration) between affected and unaffected workers according to their pre-reform earning level. Individuals earning up to 3BPCs remained unaffected by the reform and compose our control group. Instead, the treatment group comprises individuals earning an amount such that their sick leave pay would have been capped before the reform ($3BPC < 0.7w \leq 6BPC$). For these individuals, the reform increased the effective replacement rate of sick leave pay. To define treatment and control groups, we consider workers' total monthly earnings in November 2010, immediately before the reform came into force (January 2011).

13. As explained in the previous section, this is not a relevant limitation since spells fewer than 3 days old are not paid and were not affected by the reform.

We estimate the following triple difference-in-differences specification:

$$\begin{aligned} y_{it} = \alpha + \beta T_t + \gamma D_i + \delta Coop_{it} + \eta T_t \times Coop_{it} + \zeta D_i \times Coop_{it} + \theta D_i \times T_t + \\ + \phi D_i \times T_t \times Coop_{it} + \psi X_{it} + \tau_s + \omega_r + \epsilon_{it} \end{aligned} \tag{2}$$

where y_{it} either is an indicator for whether individual i experienced a sickness absence spell (lasting at least four days) in month t[14] or measures the number of days of sickness absence individual i took in month t, T_t is a post-reform dummy, D_i is the treatment group dummy, and $Coop_{it}$ is a dummy variable describing the worker cooperative status of individual i in month t. Sector τ_s and region ω_r fixed effects account for time-invariant permanent differences across 9 industries and 19 regions respectively. We also control for personal and firm-level characteristics (gender, age, tenure, firm size). Coefficient θ captures the general effect of the reform and coefficient ϕ, associated with the triple interaction, measures the differential effect for individuals employed in cooperatives. The model also includes all the corresponding two-way interactions. We estimate equation (2) by OLS, clustering standard errors at the individual level in order to account for serial correlation.

Figure 3 plots the evolution of the average sick leave benefit (in real terms) for both treatment and control individuals over time. The average sick leave pay increases in both groups. As sick leave pay is computed as a fixed fraction of the worker's total wage, this simply reflects the general increasing trend experienced by real wages in Uruguay during this period. More importantly, there is a differential increase in average sick leave pay for treatment workers starting in January 2011, suggesting that the reform hit the treatment group in the expected way. In Figure 4, we plot the fraction of workers in the treatment and control group affected by the sick pay cap before and after January 2011. As expected, the share of workers for whom the cap was binding was higher in the treatment group than in the control group before the reform, but decreased sharply after the reform.

Finally, in Figures 5 and 6, we plot the evolution of the incidence and duration of sickness absences for treated and control cooperative and conventional workers. Both figures show the evolution is similar in the pre-reform years for the four groups. Moreover, these figures reveal that treated workers employed in worker cooperatives react very differently to the sick leave reform starting in January 2011. While these figures provide preliminary visual evidence supporting the common time trend assumption, we report results from a formal event-study analysis in section 4.3.

14. If an absence spell spans over several months, the variable takes value 1 in each month.

4 Results

4.1 Exploratory analysis: fixed effects regressions

Before presenting the results of our main empirical exercise, we investigate the comparative absence behaviour of workers employed in worker cooperatives vis-à-vis employees in conventional firms by estimating a series of fixed-effects regressions. In this case, identification comes from the variability provided by workers who switch between organisational types during the period, under the assumption that sorting is driven by time-invariant characteristics. We count 1,746 switchers, which represents approximately 3% of the sample (454 workers moved from worker cooperatives to conventional firms and 1,292 made the reverse switch). The sample is restricted in the way explained in Section 3.1, except for the fact that we include all individuals regardless of their pre-reform wage.

Table 2 reports the corresponding estimates from fixed-effects regressions. We successively add controls for personal firm-level characteristics (age, tenure, firm size), year, industry, and region fixed effects. In all specifications, we cluster standard errors at the individual level. In columns (3) and (6), we report results from our preferred specifications, including region and industry-specific time trends. These estimates indicate the incidence of sickness-related absences in a given month is 1.3 percentage points higher for individuals employed in worker cooperatives compared with those employed in conventional firms. Moreover, workers employed in cooperatives spend 0.33 more days per month on sick leave compared to those employed in conventional firms. This difference is statistically significant at conventional levels.

4.2 Difference-in-differences estimates

Table 3 shows our main difference-in-differences estimates. The sample is restricted to control and treatment individuals, as defined in Section 3.2. We exploit the fact that individuals employed in worker cooperatives and conventional firms were exposed to an exogenous variation in the generosity of paid sick leave as a result of the reform. Columns 1-3 show the estimated coefficients for the incidence of sickness absences (extensive margin). In column (1) we include controls for individual- and firm-level attributes (sex, age, tenure, and firm size) and region and industry fixed effects. In column (2), we add industry- and region-specific time trends to control for time-varying shocks. In column (3), we restrict the sample to full-time workers aged 18-59.

The coefficient associated with the triple interaction term, which measures the differential effect of the reform for treated workers employed in worker cooperatives, is significantly positive in all specifications. Our estimates reported in Column (3) indicate that treated workers in cooperatives increased their probability of being absent from work in a given month by 1.6 percentage points in comparison to treated workers employed in conventional firms. This effect implies a 40% increase relative to the average pre-reform incidence of sickness absence among treated cooperative workers. Columns 4-6 report estimates considering the duration (in days) of sickness-related absences as the dependent variable. According to estimates reported in Column (6), which include industry- and region-specific time trends and restricts the sample to full-time workers aged 18-59, treated workers in cooperatives differentially increased absences by 0.4 days in a given month. The magnitude of the effect is sizeable, implying a 55% increase relative to the average pre-reform duration of sickness absence spells in that group.

4.3 Robustness checks and additional results

Event-study analysis. Our results indicate a differential intensification of absence behaviour among treated workers employed in worker cooperatives after January 2011. If the effect is due to the paid sick leave reform, we should not observe any differential pattern before 2011. Figure 7 and 8 report the results from an event-study analysis, showing the evolution of sickness-related absences over the years around the paid sick leave reform. Each estimated coefficient corresponds to the interaction between $T_t \times Coop_{i_t}$ and a full set of year dummies, where the coefficient for 2010 is normalized to zero. We do not find evidence of differential trends in workplace absences before 2011. The differential increase in sickness-related absences for treated workers employed in worker cooperatives becomes significant in 2012 and 2013.

Switchers. An important concern is that the reform may induce sorting of workers into cooperatives according to unobserved factors that may also affect their likelihood of sickness absence. We address this concern by restricting the analysis to a subsample of individuals who did not switch between conventional and worker cooperatives during this period. Our DiD estimates excluding job switchers are reported in columns 1-2 of Table 4. Treated workers in cooperatives increased their likelihood of being absent from work in a given month by 1.3 percentage points in comparison to other groups. The effect is significant at the 10% level. Duration increased by 0.316 days relative to the other groups, though the effect is not statistically significant at conventional levels (SE 0.195). This suggests that self-selection into worker cooperatives resulting from the paid sick leave reform cannot fully account for our results. Of course, we cannot rule out sorting effects in general. However, the fact that we observe a similar pre-reform trend in absence behaviour suggests that sorting pre-reform is unlikely.

Compositional changes. We perform additional DiD estimates using the balanced panel in order to control for workforce compositional changes. Estimates reported in Columns (3)-(4) of Table 4 restrict the sample to individuals observed for 24 consecutive months before and after the reform (balanced panel). We find a significant differential increase in absenteeism among treated workers in cooperatives relative to other groups. The magnitude of the effect is similar to our baseline estimates: the incidence of sickness-related absences increased by 1.5 percentage points and duration raised by 0.5 days in a given month.[15].

Individual fixed effects. We also control for time-invariant unobserved heterogeneity by estimating a difference-in-differences model with individual fixed effects. Results reported in columns (5)-(6) of Table 4 still indicate that the incidence and duration of sickness absence increased differentially among treated workers employed in cooperatives in relation to the other groups.[16]

Continuous Treatment. Our binary treatment indicator masks the fact that the increase in the generosity of paid sick leave after January 2011 did not affect individuals in the treatment group uniformly. As shown in Figure 1, the pre-reform benefit cap (3 BPC) gradually increased by 1 BPC per year from January 2011, reaching 6 BPC by January 2013. Workers earning less than 3 BPC just before January 2011 were not intended to be affected by the reform (never treated control group). Instead, all individuals earning above 3 BPC became treated in January 2011 (treatment group). Some of them, however, were also eligible to receive incremental "doses" in January 2012 and January 2013. To be more precise, the staggered intensification of the treatment worked as follows: (1) individuals earning 3-4 BPC *only* benefited from the initial sick pay cap rise in January 2011; (2) individuals earning 4-5 BPC *also* benefited from the second cap rise in January 2012; (3) finally, individuals earning 5-6 BPC were *also* eligible to benefit from an additional cap rise in January 2013. Hence, our treatment is multi-valued.

Following Ziebarth 2013, we take into account differences in treatment intensity by computing for each individual the (potential) reform-induced increase in statutory sick leave pay over the entire post-reform period relative to her pre-reform gross wage. Our measure of treatment intensity (dose) takes the value zero for workers in the control group and positive values up to 35% of workers' gross wage for those in the treatment group. On average, the potential sick leave benefit for treated workers increased by 19% of their gross wage due to the reform. Results are presented in columns (7)-(8) of Table 4. Consistent with our previous results using a discrete treatment indicator, the behavioural response to treatment intensity for workers employed in cooperatives is

15. We also estimate a more flexible DiD model interacting individual (gender, age, tenure) and firm-level characteristics (size, region, industry) with our Post-reform, treatment, and worker cooperative dummies. This model allows covariates to have a differential effect depending on time and individuals' treatment and cooperative status. Reassuringly, results are very similar to our baseline estimates (see Appendix Table A.1.1)

16. It is worth noting that in this case the effect is identified from within-individual change in their $D \times T$ and $D \times T \times Coop$ status over time.

significantly stronger relative to other groups.

In a recent paper, Callaway, Goodman-Bacon, and Sant'Anna 2021 identify crucial weaknesses of the DiD estimator in the presence of multiple time periods and variation in treatment intensity and timing of adoption. In particular, they identify a bias arising from a specific form of selection into different amounts of the treatment (selection-on-gains). Hence, to compare treatment effects across groups exposed to different dosage levels would require to make stronger assumptions than in the standard DiD framework. Unfortunately, there is no straightforward practical solution to this problem in the literature. As mentioned, our treatment group is composed of individuals who only benefited from the initial sick pay cap hike and individuals who also benefited from subsequent cap rises in January 2012 and January 2013. Therefore, we should assume that the treatment effect for the first group is the same as the treatment effect for the other group had they both benefited only from the first benefit cap rise. This would be violated if individuals have the discretion to set their pre-reform wages and, hence, dosage level based on their expected gains from the reform.

To further dig into this issue, we estimate a separate DiD model comparing individuals in the control group with individuals who experienced the same treatment intensity and timing. We focus on the group of individuals earning 3-4 BPCs just before January 2011, who only benefited from the first sick pay cap hike. Results reported in Appendix Table A.1.2 are qualitatively similar to our baseline estimates.

5 Mechanisms

Short-term vs. long-term absenteeism. The Uruguayan sick leave insurance system does not make any distinction between short- and long-term absences in terms of replacement rates and funding. However, the distinction might be important to understand the underlying mechanisms behind the differential response of individuals employed in worker cooperatives. Assuming that individuals on long-term sick leave are more prone to be seriously sick, it has been argued that standard labour supply responses driven by moral hazard might be more relevant for short-term rather than for long-term sickness absence. Following Ziebarth 2013, in a given month, we classify sickness-related absences originated in absence spells lasting more than 6 weeks as long-term absences. In our sample, long-term absences account for 53% of all absence days although they only represent 21% of all sickness cases.

In columns (1) and (2) of Table 5, we display estimates of equation (2) of the incidence of sickness absence for short- and long-term sickness spells, respectively. Results reported in column (1) indicate that the incidence of short-term absences for treated individuals employed in worker cooperatives increased by 0.4 percentage points relative to treated workers employed in conventional firm. We also find that long-term absences among individuals employed in cooperatives increased by 1 percentage points relative to the other group. The change in short-term absences suggests that the increase in workplace absenteeism in cooperatives after the reform is partly attributable to moral hazard problems. Interestingly, cooperatives also seem to facilitate greater take-up of long-term sick leave, presumably motivated by genuine health conditions.

Members vs. employees in worker cooperatives. As explained in section 2.2, worker cooperatives can also hire employees at market wages as do conventional firms. The distinction between members and employees in worker cooperatives is relevant in our context given the different incentive structure faced by the two types of workers, which in turn may affect their responses to the paid sick leave reform. In contrast to members, hired workers in cooperatives do not participate in strategic managerial decisions and do not have an ownership stake in the firm. Therefore, one could hypothesise that members and hired employees in worker cooperatives face different labour discipline environments. For instance, the threat of dismissal due to unsatisfactory job performance may be less credible in the case of members.[17]

In columns (3) and (5) of Table 5, we report DiD estimates comparing individuals employed in conventional firms and members of worker cooperatives, while columns (4) and (6) display estimates only comparing employees in conventional firms and worker cooperatives. Interestingly, the differential behavioural response of affected individual employed in worker cooperatives in terms of both incidence and duration of absence spells is entirely driven by the behaviour of cooperative members.

Extensive margin responses by disease-categories. In this section, we further investigate extensive margin responses of sickness absence to the paid sick leave reform by exploiting information on doctor-certified disease categories. Using medical diagnosis classified according to the International Classification of Diseases (ICD), we analyse six broad categories of diseases: musculoskeletal, infectious, respiratory[18] , mental, poisoning, and pregnancy complications. The anatomy of responses by certified disease categories may be informative of the underlying mechanisms behind individuals' behavioural responses in worker cooperatives. In particular, the comparison between labour supply adjustments for musculoskeletal (e.g. back pain) and infectious diseases has proved helpful in unpacking responses to paid sick leave in terms of shirking behaviour and contagious presenteeism (Pichler and Ziebarth 2017).

17. Interviews with managers of the world's biggest (and recently demised) industrial worker cooperative indicate that members' absenteeism was an important concern: *"The moment they became members, their sense of commitment just slipped away.(. . .) Being a member was almost like being in the public service. Absenteeism skyrocketed, especially on Mondays. I think it was a lack of commitment. And I think Human Resources should have come down harder on them"*(Basterretxea, Heras-Saizarbitoria, and Lertxundi 2019, p.592).

18. Respiratory diseases are part of a mixed category including both contagious and noncontagious diseases.

Information on disease categories for each sickness spell is only available from 2010 onward. Hence, we redefine our treatment and control groups and compare the pre-reform (2011-2012) and post-reform period (2013), exploiting the increase in the benefit cap that came into force in January 2013. Table 6 displays our disease-specific DiD estimates. We observe a differential increase in the incidence of musculoskeletal conditions for treated individuals in worker cooperatives relative to the other group. This category includes hard-to-diagnose conditions (e.g. back pain) and is more prone to moral hazard reporting problems.

Marginal utility of leisure: extended weekends absences. We further exploit the granularity of the data to see whether sickness absences in cooperatives are more frequent on days in which leisure may confer greater marginal utility. A crucial advantage of the data is that we know the precise start and end date of each sickness spell.[19]

As a first approximation, we investigate the existence of a "Monday effect." Figure 9 plots the distribution of sickness spells by day of first report.[20] If the start of a sickness spell is randomly distributed over the week, one should expect 20% of them to start on Monday. We observe that an excess proportion (5 percentage points) of spells started on Mondays. The pattern appears to be very similar for individuals employed in cooperatives and conventional firms. In Table 7, we report additional DiD estimates of the incidence of sickness spells by the day of first report. There are individuals with multiple absence spells in a given month. For this reason, estimates consider the day of first report of each absence spell in a given month. There is no evidence of a differential increase in extended weekend absences (Monday/Friday) for treated individuals employed in cooperatives compared to the other groups.

Labour discipline. We also investigate whether documented differences in absence behaviour between individuals employed in cooperatives and conventional firms could be explained by the use of more punitive labour discipline strategies in conventional firms. It is a well-established fact that worker cooperatives have more stable employment and destroy fewer jobs than conventional firms (Craig and Pencavel 1992; Pencavel, Pistaferri, and Schivardi 2006; Burdín and Dean 2009; Alves, Burdín, and Dean 2016). Union members are also less likely to lose their jobs than non-members, which, in turn, may explain why they react more strongly to variations in paid sick leave (Goerke and Pannenberg 2011, 2015). It is natural to think that a similar mechanism could be at work when employees have full bargaining power as in worker cooperatives.

19. The existence of the so-called "Monday effect" has been studied in the context of U.S. workers' compensation programs providing insurance against work-related injuries (Card and McCall 1996; Campolieti and Hyatt 2006). Related papers have analysed the impact of pubic holidays, weather conditions, sport events, and birthdays on absence behaviour (Böheim and Leoni 2019, Shi and Skuterud 2015, Thoursie 2004, Thoursie 2007).

20. In Appendix Figure A.1.1 and A.1.2, we report the distribution of sickness spells by disease categories and day of first report.

We adopt a similar DiD approach, comparing the evolution of dismissal rates between treated and control workers in both types of firms before and after the increase in sick leave pay. We identified dismissed individuals in each month by relying on both administrative information on the cause of separation (i.e. dismissal) and whether the individual was receiving unemployment benefits. In this way, we are able to restrict the analysis to layoffs, excluding other types of separations (voluntary terminations, retirement, etc.).

Table 9 shows estimates of equation (2) in which the dependent variable is a dummy indicating that a worker has been fired in the respective month. Our preferred estimates reported in column (2) indicate that the probability of being individually dismissed is 0.8 percentage points lower among treated workers employed in worker cooperatives relative to other groups. Considering the average pre-reform dismissal rate (1%), the magnitude of the effect is large. In column (3)-(5), we show that differences in the use of layoffs are explained by individuals employed in large firms and are larger for worker-members.

Figure 10 reports the results from an event-study analysis in which we track differences in dismissal rates before and after the paid sick leave reform. As in column (5) of 9, we consider individuals employed in conventional firms and members of worker cooperatives. Each estimated coefficient corresponds to the interaction between $T_t \times ConventionalFirm_{it}$ and a full set of year dummies, where the coefficient for 2010 is normalized to zero. The differential increase in dismissal rates for treated workers employed in conventional firms relative to cooperatives becomes positive and significant from 2011 onward. We observe broadly similar trends in the likelihood of dismissal before the reform, although there is a statistically significant violation of parallel pre-trends in 2008. Our analysis of the dynamics of layoffs is at least suggestive that conventional firms relied on more punitive labour discipline strategies than did cooperatives and were more prone to use the threat of dismissal after the reform.

Small vs. large firms. It has been argued that cooperative teams and profit sharing arrangements may suffer from weak work incentives (Alchian and Demsetz 1972). However, the extent of free riding may vary with the size of the team. Large teams may be particularly vulnerable to shirking behaviour (*1/n problem*). By contrast, in small teams, the dilution of incentives may be less severe and shrinking could be mitigated through mutual monitoring among members without relying on specialised supervisors. To check for this mechanism, in Table 8 we present additional estimates splitting the sample by firm size. We define small firms as those with less than 20 workers. The differential increase of absenteeism in cooperatives holds only for individuals employed in medium-sized and large firms, though differences appear to be larger in the subsample of large firms. In the next section, we report additional survey evidence on managers' perceptions about work ethics in both types of firms, confirming that large cooperatives were particularly affected by absenteeism after the reform.[21]

6 Worker supervision and managers' perceptions of work ethics: additional survey evidence

To further sharpen the interpretation of our findings, we provide additional firm-level survey evidence on managers' perceptions about absenteeism and work ethics collected before (2009) and after (2012) the reform. We collected information on a sample of roughly 400 Uruguayan firms per wave, including both worker cooperatives and conventional firms. By design, the comparison group of conventional firms mimics the sectoral and size distribution of cooperatives. In what follows, we restrict the analysis to firms that responded to the survey in both waves.

In Appendix Figure A.1.3 (Panel A), we report managers' responses to the following question: *¿Could you rank the most pressing human resource management problems faced by your company during the last year?* The evidence suggests that absenteeism is perceived as the main HRM problem in medium-large cooperatives. Moreover, concerns about absenteeism among managers of large cooperatives increased sharply between survey waves, coinciding with the implementation of the reform. In Figure A.1.3 (Panel B), we report responses to the following question: *What is your perception of work attitudes that predominate among individuals employed in your company?* This question is only available for the post-reform wave. The share of managers perceiving a low or very low work ethics in their companies is larger among managers of medium-large cooperatives. By contrast, poor work ethics does not seem to be a problem for small cooperatives. Altogether, survey evidence appears to be broadly consistent with our DiD estimates by firm size.

To further understand the distinct labour discipline environment of worker cooperatives, we report information on supervision intensity and monitoring mechanisms. In Figure A.1.4 (Panel A), we display the supervision intensity by firm size and organisational form. We define the supervision ratio as the number of supervisors divided by total employment.[22] In the case of small firms, supervision intensity appears to be lower in worker cooperatives than in conventional firms. By contrast, large firms

21. In Appendix A.1.1, we present a complementary empirical exercise comparing individuals' absence behaviour before and after a worker buyout, i.e. the conversion of a conventional firm into a worker cooperative. We distinguish worker buyouts of small and large firms. Interestingly, we only observe a significant increase in absenteeism after a worker buyout of a large firm.

22. Specifically, the questionnaire asks managers to report the number of workers performing supervision tasks. Following Wright 1995 and Jayadev and Bowles 2006, supervisors are defined as workers that have more than one subordinate and can make decisions regarding the tasks, the tools or procedures to be used, and the pace of work of their subordinates. They can also sanction (or cause to be sanctioned) with respect to pay, promotions or job termination.

exhibit roughly similar supervision ratios, regardless of their organisational form. Finally, we ask managers to report the main mechanism used by the firm to monitor and enforce work effort. In Figure A.1.4 (Panel B and C), we show that hierarchical monitoring by specialized supervisors (*"Verbal warnings from supervisors"*) is more common in conventional firms, while mutual monitoring among coworkers (*"Verbal warnings from coworkers"*) is more frequent in cooperatives. Interestingly, despite exhibiting a roughly similar supervision intensity, supervisors in medium-large cooperatives are perceived as less active in enforcing labour discipline than supervisors in conventional firms of similar size. Surprisingly, peer monitoring is also a relevant disciplinary mechanisms among medium-sized and large cooperatives. However, the documented differences in absence behaviour and perceived work ethics between cooperatives of different sizes suggest that peer monitoring constitutes a feasible substitute for hierarchical supervision only in the context of small cooperatives.

7 Conclusions

In this paper we aim at understanding individuals' effort choices, proxied by absence behaviour, across different organisational settings. Using monthly employment history data matched with individual-level sick leave records and exploiting an exogenous increase in the paid sick leave maximum cap in Uruguay, we compare the absence behaviour of individuals employed in worker cooperatives and in conventional firms. A worker cooperative constitutes a rather peculiar organisational setting in which worker-members have a stake in ownership and ultimately controlled managerial decisions.

We find a differential increase in absence behaviour among treated individuals employed in a worker cooperative relative to individuals employed in conventional firms. Differences between the two groups are driven by both short-term and long-term absences, members' behaviour, hard-to-diagnose conditions, and individuals employed in medium-sized and large cooperatives. We also find suggestive evidence that, relative to worker cooperatives, conventional firms employ dismissals more frequently as a disciplinary tool to reduce absenteeism after the reform. Small cooperatives did not suffer from a similar increase in absenteeism. Altogether, our findings indicate that conventional effort supply responses driven by moral hazard account for at least part of the differential increase in absenteeism among workers in cooperatives. Survey evidence on managers' perceptions suggests lower perceived work ethics in large cooperatives, where peer monitoring may be less feasible as an alternative labour discipline device.

The social welfare implications of individuals' behaviour under the two organisational settings are not straightforward. On the one hand, our findings suggest that a potential non-pecuniary benefit from cooperative membership could be a more discretionary utilisation of voluntary absences. This may come at a cost in terms of firm output, particularly in the context of large cooperative teams. On the other hand, conventional firms require the use of layoffs to enforce labour discipline and keep absenteeism under control. This entails potential negative externalities as firms do not fully internalise the consequences of layoffs for individual welfare and public finances. Moreover, workers may underutilize sick leave insurance, leading to potential problems of contagious presenteeism, reduced productivity, and additional costs to public health services. Further research could analyze how differences in absence behaviour map into productivity gaps between the two types of firms. The answer is not obvious as organisations may differ in their ability to replace absent workers and avoid disruptions in the production process.

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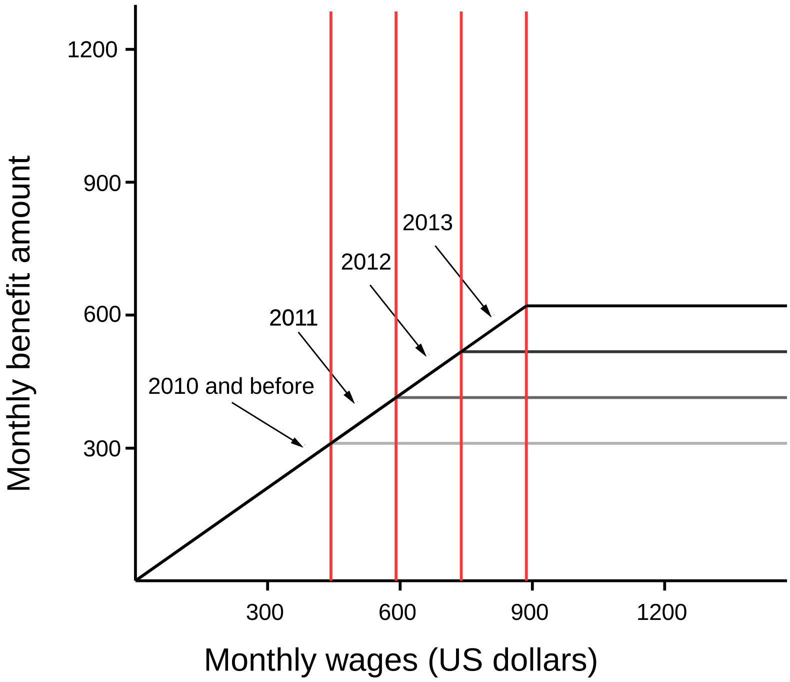
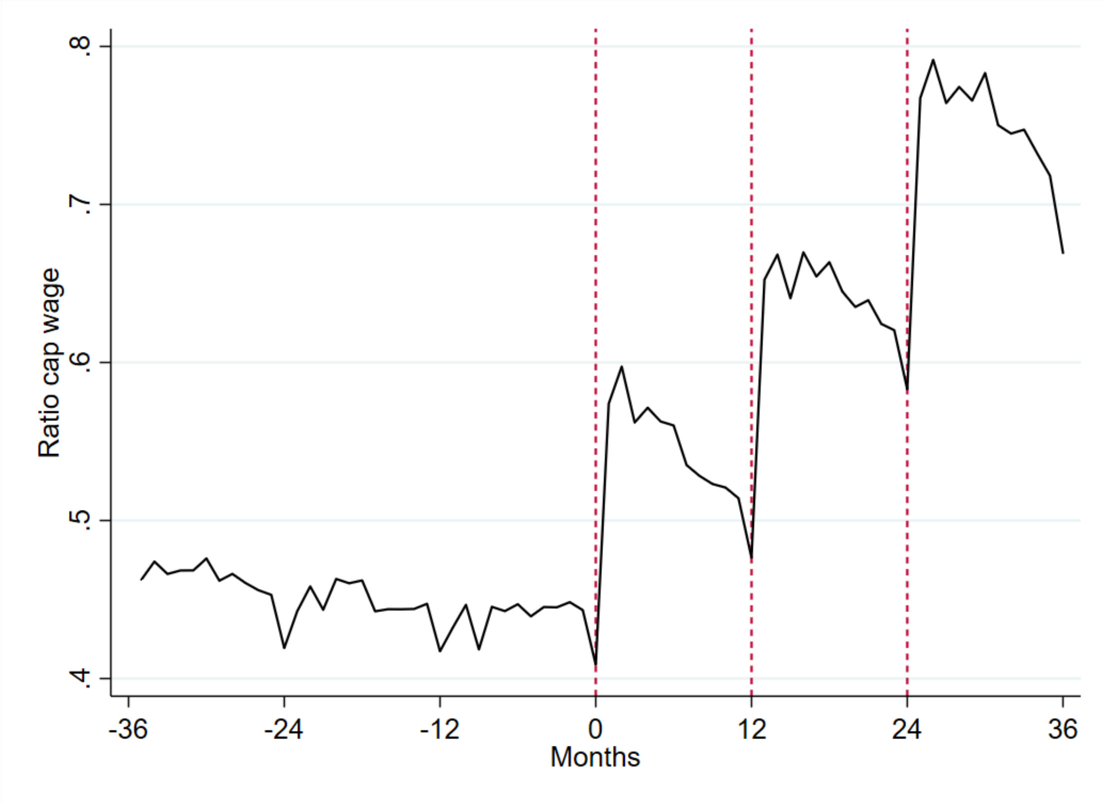
Figures and Tables

Figure 1: Paid sick leave schedule before and after the reform



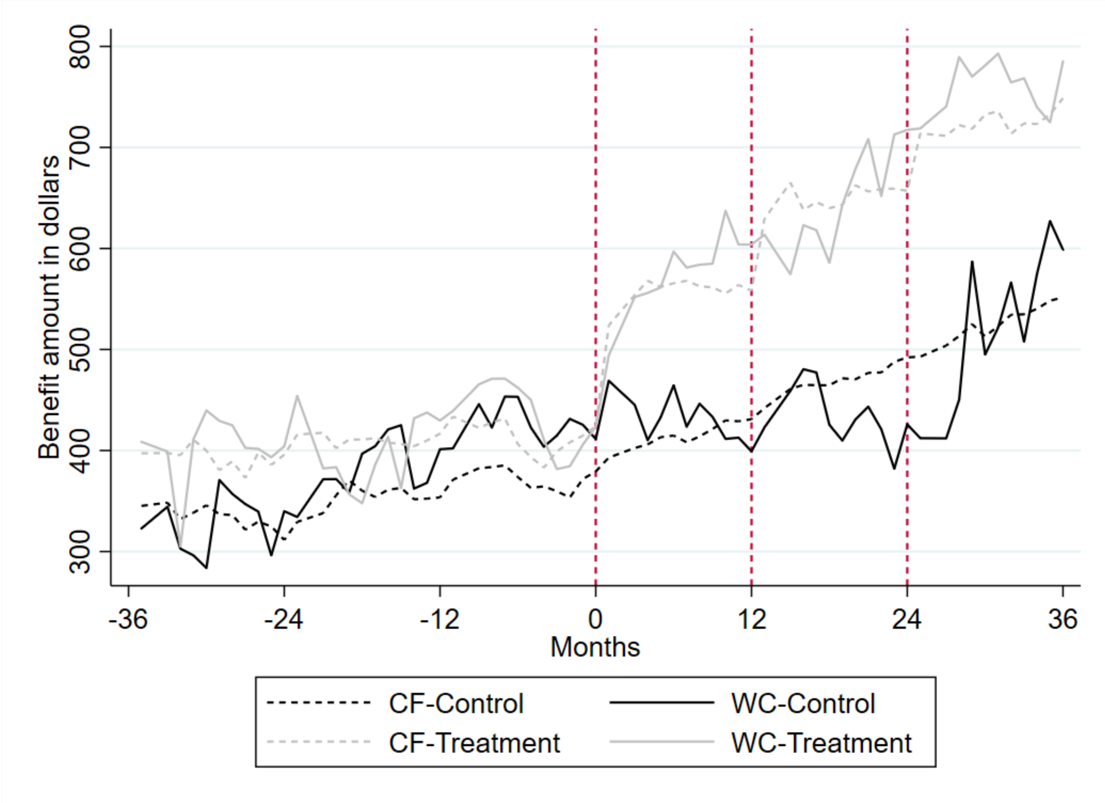
Notes: Authors' elaboration based on provisions of the Sick Leave Insurance Law 18725 (December 2010). The graph shows the evolution of the schedule of the paid sick leave monthly benefit amount in nominal terms (USD) as a kinked function of previous earnings in Uruguay. Changes in the maximum benefit amount also apply to the benefit amount of ongoing spells.

Figure 2: Ratio between paid sick leave benefit cap and average wage



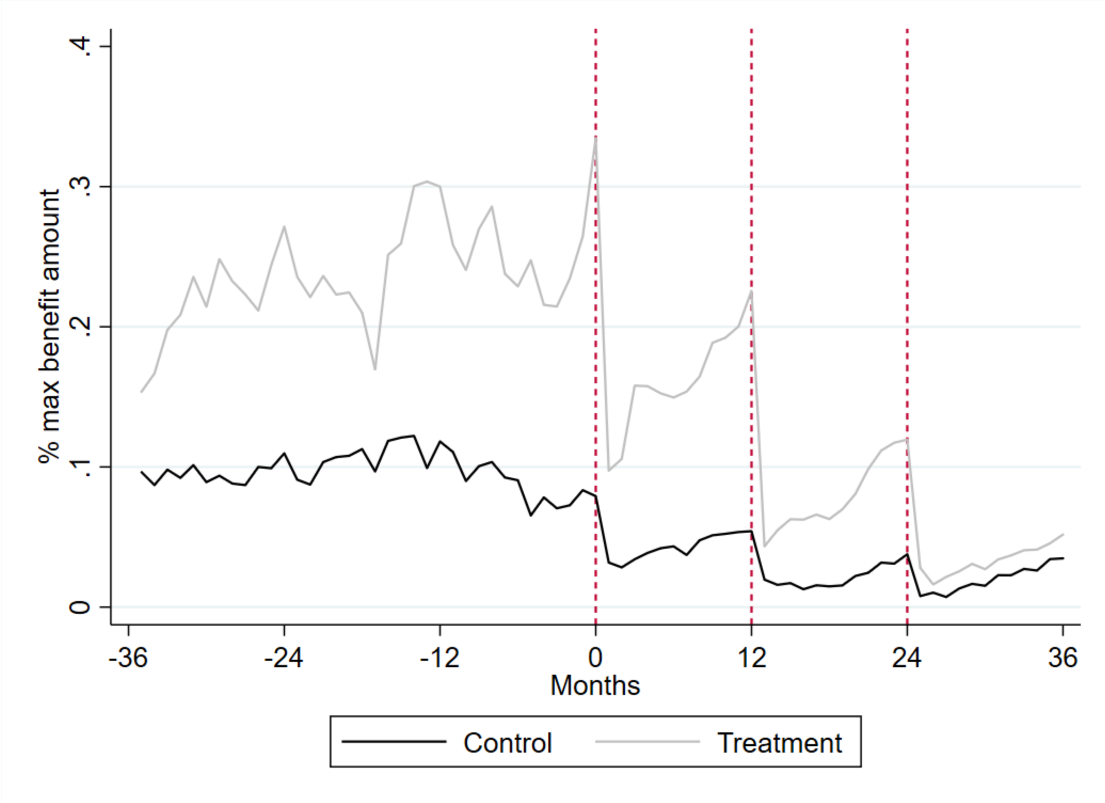
Notes: The graph shows the evolution of the sick pay cap relative to average wages.

Figure 3: Evolution of average paid sick leave by treatment status and organisational form



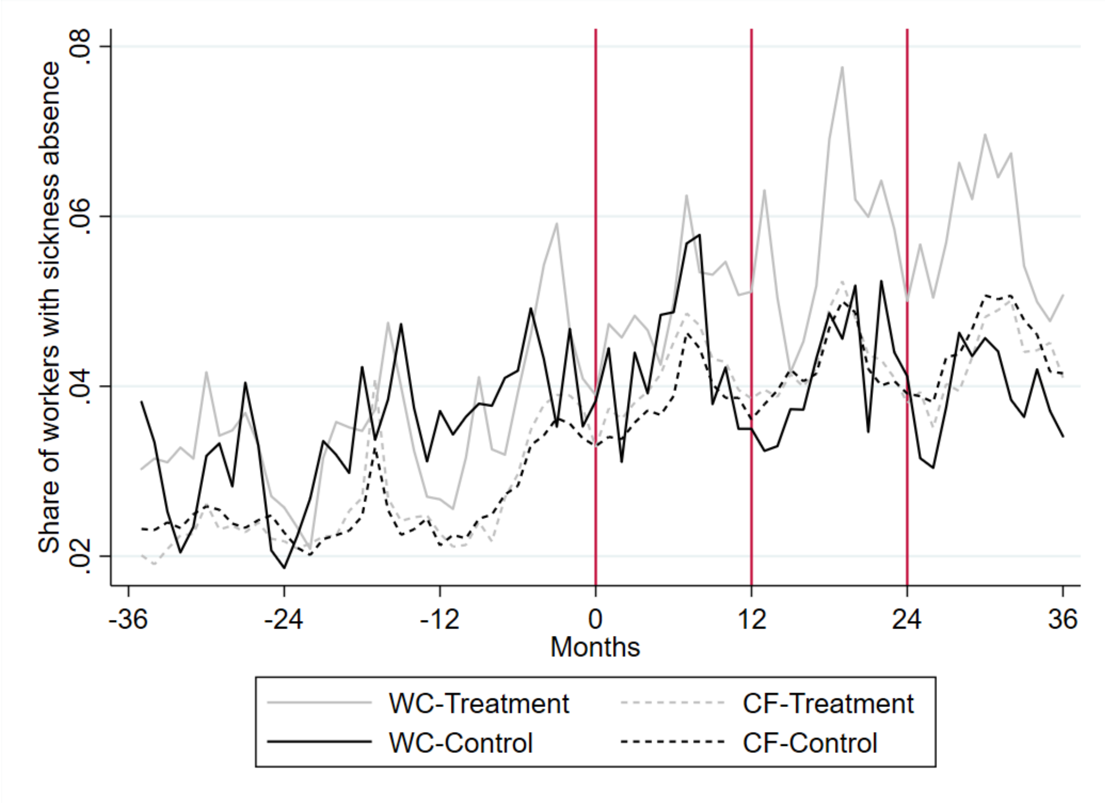
Notes: The graph displays the evolution of average sick pay for treatment and control groups in conventional firms (CF) and worker cooperatives (WC) before and after the reform.

Figure 4: Ratio of workers affected by the benefit cap



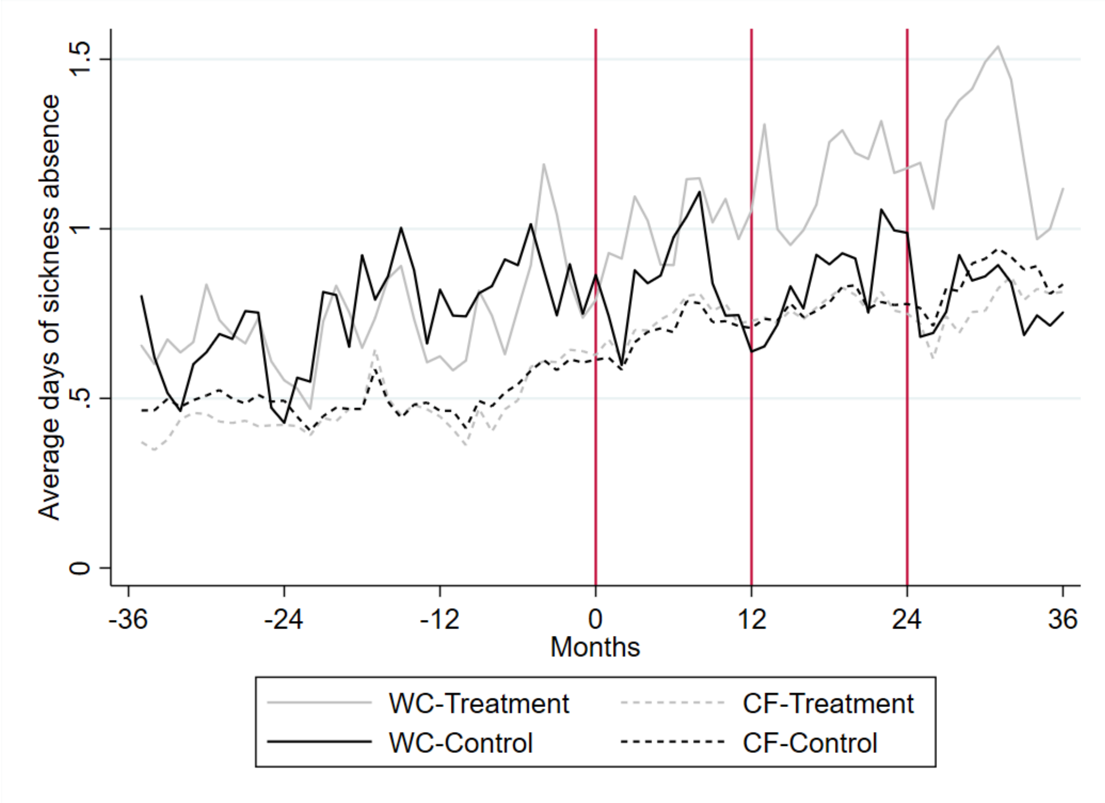
Notes:The graph displays the share of individuals affected by the sick pay cap in treatment and control groups before and after the reform.

Figure 5: Share of workers with sickness absence in each month



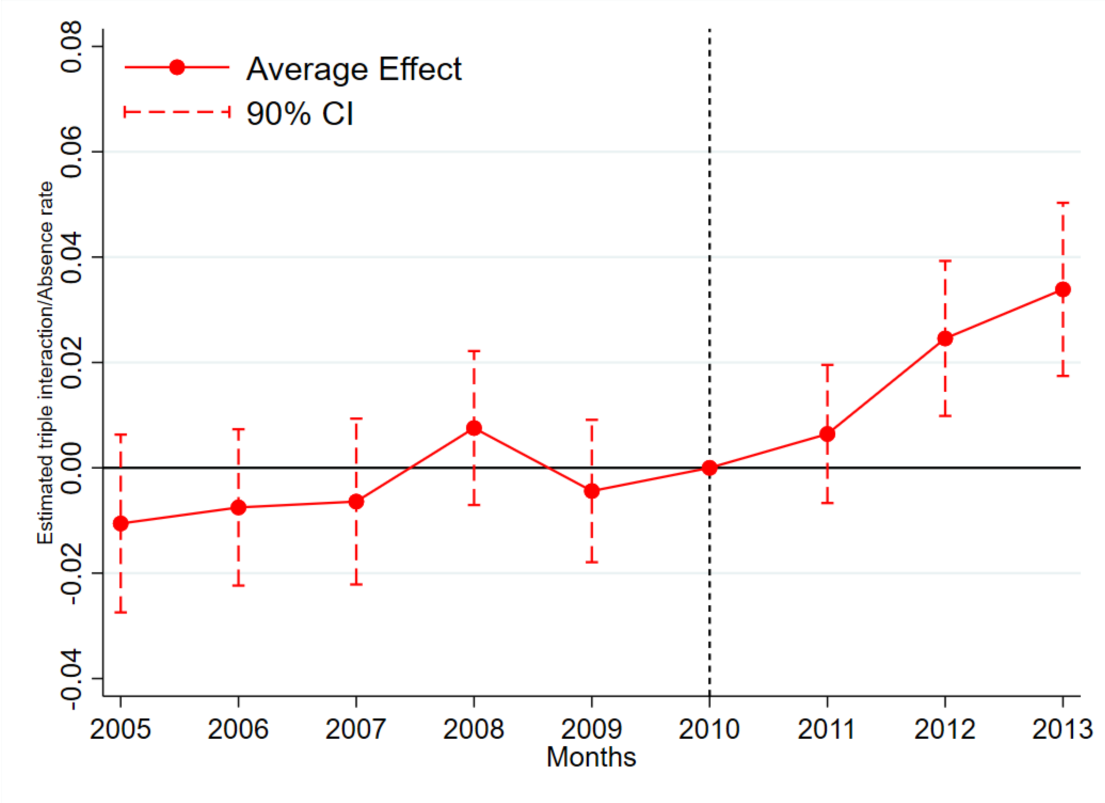
Notes: The graph displays the share of individuals with a sickness-related absence (lasting at least four days) in each month. The figure distinguishes treatment and controls in conventional firms (CF) and worker cooperatives (WC) before and after the reform.

Figure 6: Average duration of sickness absence spells (in days)



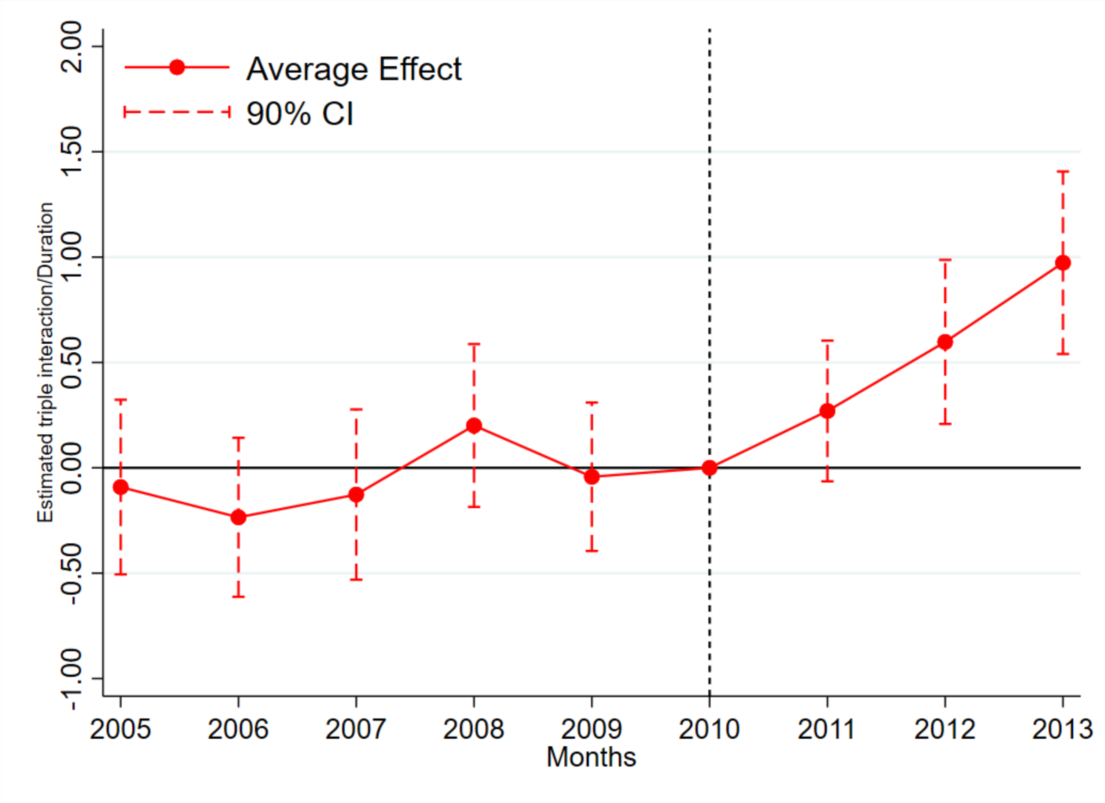
Notes: The graph displays the average duration (in days) of sickness-related absence spells in each month. The figure distinguishes treatment and controls in conventional firms (CF) and worker cooperatives (WC) before and after the reform.

Figure 7: Event-study analysis: incidence of sickness-related absence



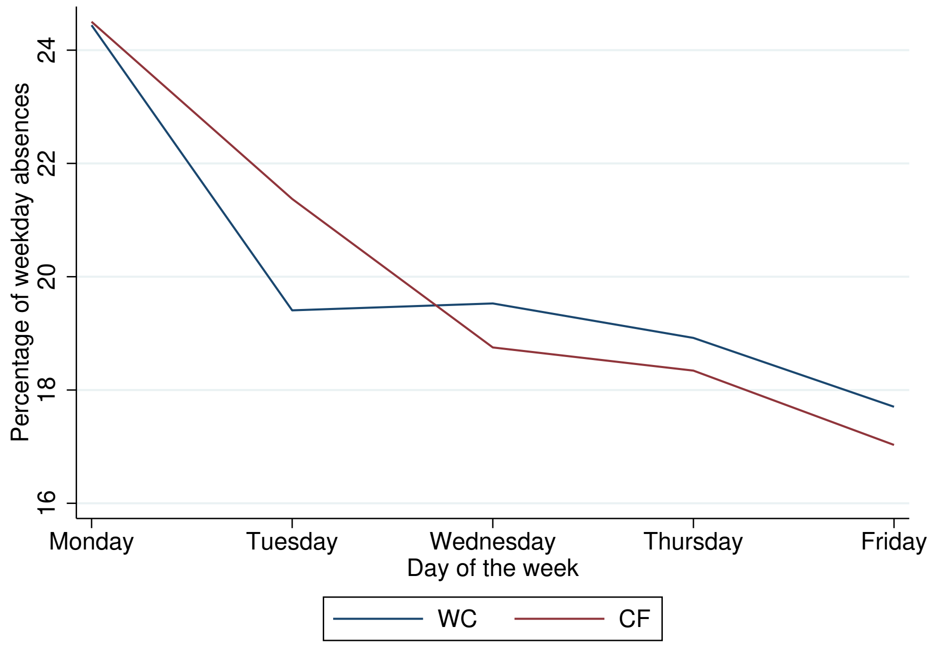
Notes: The figure shows event studies based on a triple DiD model as in Equation (2). Dependent variable: indicator for whether individual i experienced a sickness absence spell (lasting at least four days) in month t. The graph displays the estimated ϕ coefficient associated with the triple interaction term $D_i \times T_t \times Coop_{it}$, i.e. the heterogeneous effect by organisational form (employees in conventional firms vs. members in worker cooperatives). The standard errors are clustered at the individual level and the dash bars depict 90% confidence intervals.

Figure 8: Event-study analysis: duration of sickness-related absence



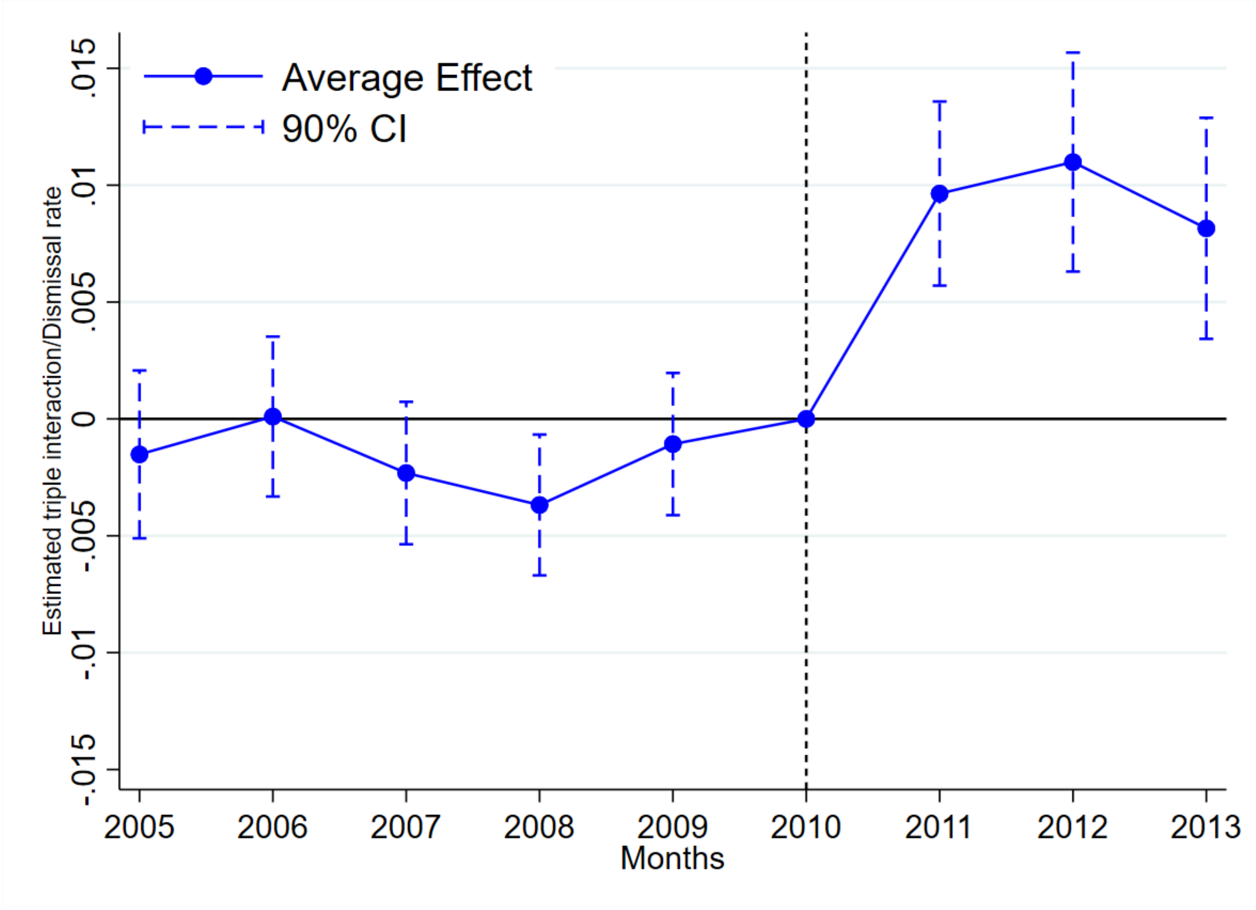
Notes: The figure shows event studies based on a triple DiD model as in Equation (2). Dependent variable: number of days of sickness absence individual i took in month t. The graph displays the estimated ϕ coefficient associated with the triple interaction term $D_i \times T_t \times Coop_{it}$, i.e. the heterogeneous effect by organisational form (employees in conventional firms vs. members in worker cooperatives). The standard errors are clustered at the individual level and the dash bars depict 90% confidence intervals.

Figure 9: Distribution of sickness-related absence spells by day of first report



Notes: The figure displays the distribution of sickness-absence spells by day of first report, distinguishing individuals employed in conventional firms (CF) and worker cooperatives (WC).

Figure 10: Event-study analysis: dismissals



Notes: The figure shows event studies based on a triple DiD model as in Equation (2), but substituting the dummy $Coop_{it}$ for the dummy $ConventionalFirm_{it}$ that takes the value 1 for conventional firms. Dependent variable: indicator for whether individual i experienced a layoff in month t. The graph displays the estimated ϕ coefficient associated with the triple interaction term $D_i \times T_t \times ConventionalFirm_{it}$, i.e. the heterogeneous effect by organisational form (employees in conventional firms vs. members in worker cooperatives). The standard errors are clustered at the individual level and the dash bars depict 90% confidence intervals.

Table 1: Descriptive statistics

	Pre-reform (2008-2010)				Post-reform (2011-2013)			
	Control		Treatment		Control		Treatment	
	CFs	Coops	CFs	Coops	CFs	Coops	CFs	Coops
Incidence of sickness-related absences (monthly)	0.03	0.03	0.03	0.04	0.04	0.04	0.04	0.06
Duration of absence spells (days)	0.50	0.75	0.47	0.73	0.77	0.83	0.76	1.15
Age	36.28	41.45	37.06	46.15	37.62	42.56	39.21	46.86
% Male	0.47	0.44	0.63	0.79	0.48	0.46	0.63	0.75
Tenure (years)	2.97	3.94	4.47	4.96	3.64	4.56	5.62	6.36
Average salary of the firm (log)	2.25	2.04	2.87	2.75	2.62	2.42	3.16	3.01
Number of workers (log)	2.76	3.24	3.73	3.46	2.97	3.53	3.83	3.51
% Part-time worker	0.22	0.47	0.07	0.12	0.22	0.50	0.08	0.16
% Small firms	0.84	0.74	0.72	0.73	0.82	0.71	0.69	0.67
% Manufacturing	0.16	0.16	0.18	0.07	0.15	0.15	0.17	0.08
Average observations by month	18,888	700	16,446	987	20,377	674	15,652	942

Notes: Authors' elaboration based on monthly employment administrative records. Uruguayan Social Security Agency (Banco de Prevision Social).

Table 2: Incidence and duration of sickness-related absence: fixed-effects regressions

	Incidence of sickness-related absence			Duration (days)		
	(1)	(2)	(3)	(4)	(5)	(6)
Coop	0.017***	0.013**	0.013**	0.354***	0.345**	0.325**
	(0.005)	(0.006)	(0.006)	(0.128)	(0.139)	(0.139)
Observations	2,987,831	2,644,898	2,644,898	2,987,831	2,644,898	2,644,898
R-squared	0.161	0.162	0.162	0.199	0.197	0.198
Individual's controls	No	Yes	Yes	No	Yes	Yes
Industry-specific time trends	No	No	Yes	No	No	Yes
Region-specific time trends	No	No	Yes	No	No	Yes

Notes: Individual-level controls include age, male, tenure, firm size (log of total employment), 9 industry dummies, 19 regional dummies ("Departamentos"). Standard errors clustered at the individual level are reported in parentheses. *** p<0.01, ** p<0.05, * p<0.1

Table 3: Difference-in-differences estimates

	Incidence of sickness-related absence			Duration (days)		
	(1)	(2)	(3)	(4)	(5)	(6)
Post-Reform × Treatment	0.002*	0.002*	0.001	0.036	0.035	0.023
	(0.001)	(0.001)	(0.001)	(0.027)	(0.028)	(0.032)
Post-Reform × Treatment × Coop	0.011*	0.014**	0.016**	0.292*	0.357**	0.415**
	(0.006)	(0.006)	(0.008)	(0.151)	(0.154)	(0.208)
Observations	2,395,433	2,395,433	1,719,958	2,395,433	2,395,433	1,719,958
R-squared	0.019	0.020	0.018	0.015	0.015	0.015
Individual's controls	Yes	Yes	Yes	Yes	Yes	Yes
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes
Industry-specific time trends	No	Yes	Yes	No	Yes	Yes
Region-specific time trends	No	Yes	Yes	No	Yes	Yes
Only full-time workers aged 18-59 years	No	No	Yes	No	No	Yes

Notes: DiD estimates comparing treatment and control individuals. Estimates reported in columns 3 and 6 are restricted to workers aged 25-55 years old and employed full time. Coop equals 1 for individuals employed in a worker cooperative in a particular month and 0 otherwise. The post-reform variable equals 1 for years 2011-2013 (policy-on period) and 0 for years 2008-2010 (policy-off period). Individual-level controls include age, male, tenure, firm size (log of total employment), 9 industry dummies, 19 regional dummies ("Departamentos"). Standard errors clustered at the individual level are reported in parentheses. *** $p<0.01$, ** $p<0.05$, * $p<0.1$

Table 4: Robustness checks: switchers, balanced panel, individual FE, and treatment intensity

	Excluding switchers		Balanced panel		Individual Fixed Effects		Treatment intensity	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Incidence	Duration	Incidence	Duration	Incidence	Duration	Incidence	Duration
Post-Reform × Treatment	0.002*	0.030	0.003**	0.054	-0.001	-0.058**		
	(0.001)	(0.028)	(0.001)	(0.037)	(0.001)	(0.029)		
Post-Reform × Treatment × Coop	0.013*	0.316	0.015*	0.462**	0.013**	0.325**		
	(0.007)	(0.195)	(0.008)	(0.223)	(0.006)	(0.162)		
Post-Reform × Treatment Intensity							0.005	0.087
							(0.005)	(0.119)
Post-Reform × Treatment Intensity × Coop							0.054**	1.376**
							(0.022)	(0.596)
Observations	2,269,160	2,269,160	915,511	915,511	2,395,433	2,395,433	2,395,433	2,395,433
R-squared	0.019	0.015	0.019	0.015	0.191	0.238	0.020	0.015
Individual's controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year dummies	Yes	Yes						
Industry-specific time trends	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Region-specific time trends	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: DiD estimates comparing treatment and control individuals. In columns 1-2, we report estimates excluding individuals who switched between worker cooperatives and conventional firms. In columns 3-4, we restrict the analysis to the balanced panel (individuals with continuous work history 24 month before-after January 2011). In column 5-6, we report estimates including individual fixed effects. In column 7-8, we report estimates using a treatment intensity indicator instead of a binary one. Coop equals 1 for individuals employed in a worker cooperative in a particular month and 0 otherwise. The post-reform variable equals 1 for years 2011-2013 (policy-on period) and 0 for years 2008-2010 (policy-off period). Individual-level controls include age, male, tenure, firm size (log of total employment), 9 industry dummies, 19 regional dummies ("Departamentos"). Standard errors clustered at the individual level are reported in parentheses. *** $p<0.01$, ** $p<0.05$, * $p<0.1$

Table 5: Heterogeneous effects and mechanisms: short-term vs. long-term absences, members vs. employees

	Incidence of sickness-related absence				Duration (days)	
	(1)	(2)	(3)	(4)	(5)	(6)
	Short-term absences	Long-term absences (>6 weeks)	Only members in worker coops	Only hired workers in worker coops	Only members in worker coops	Only hired workers in worker coops
Post-Reform × Treatment	0.001	0.001	0.001	0.004***	0.013	0.013
	(0.000)	(0.001)	(0.001)	(0.001)	(0.030)	(0.030)
Post-Reform × Treatment × Coop	0.004*	0.010*	0.020***	-0.009	0.552***	-0.356
	(0.002)	(0.005)	(0.006)	(0.013)	(0.167)	(0.344)
Observations	2,395,433	2,395,433	2,159,708	2,056,824	2,159,708	2,083,876
R-squared	0.010	0.011	0.018	0.026	0.014	0.014
Individual's controls	Yes	Yes	Yes	Yes	Yes	Yes
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes
Industry-specific time trends	Yes	Yes	Yes	Yes	Yes	Yes
Region-specific time trends	Yes	Yes	Yes	Yes	Yes	Yes

Notes: DiD estimates comparing treatment and control individuals. In columns 1-2, we report estimates considering short-term and long-term absences, respectively. In columns 3 and 5, we restrict the analysis to employees in conventional firms and members of worker cooperatives. In columns 4 and 6, we restrict our DiD estimates to employees in both types of firms. Coop equals 1 for individuals employed in a worker cooperative in a particular month and 0 otherwise. The post-reform variable equals 1 for years 2011-2013 (policy-on period) and 0 for years 2008-2010 (policy-off period). Individual-level controls include age, male, tenure, firm size (log of total employment), 9 industry dummies, 19 regional dummies ("Departamentos"). Standard errors clustered at the individual level are reported in parentheses. *** $p<0.01$, ** $p<0.05$, * $p<0.1$

Table 6: Difference-in-differences estimates: incidence of sickness absence by disease categories

	(1) Musculoskeletal	(2) Infectious	(3) Respiratory	(4) Mental	(5) Poisoning	(6) Pregnancy complications
Post-Reform × Treatment	0.000	0.000	0.001**	0.000	-0.000	0.002
	(0.001)	(0.000)	(0.000)	(0.001)	(0.001)	(0.002)
Post-Reform × Treatment × Coop	0.009**	0.000	0.002	0.004	0.004	-0.001
	(0.004)	(0.001)	(0.002)	(0.003)	(0.003)	(0.006)
Observations	853,293	847,206	849,994	848,461	849,816	261,784
R-squared	0.007	0.001	0.002	0.004	0.001	0.005
Individual's controls	Yes	Yes	Yes	Yes	Yes	Yes
Industry-specific time trends	Yes	Yes	Yes	Yes	Yes	Yes
Region-specific time trends	Yes	Yes	Yes	Yes	Yes	Yes

Notes: DiD estimates comparing treatment and control individuals. Estimates reported in columns 6 are restricted to female workers. Coop equals 1 for individuals employed in a worker cooperative in a particular month and 0 otherwise. The post-reform variable equals 1 for 2013 (policy-on period) and 0 for years 2011-2012 (policy-off period). Individual-level controls include age, male, tenure, firm size (log of total employment), 9 industry dummies, 19 regional dummies ("Departamentos"). Standard errors clustered at the individual level are reported in parentheses. *** p<0.01, ** p<0.05, * p<0.1

Table 7: Difference-in-differences estimates: day of first report

	(1) All	(2) Monday	(3) Tuesday	(4) Wednesday	(5) Thursday	(6) Friday	(7) Monday-Friday	(8) Tuesday-Thursday
Post-Reform x Treatment	0.001**	0.000	0.000	0.000	0.000*	0.000	0.001***	0.001***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Post-Reform x Treatment x Coop	0.003*	0.001	0.000	0.001	-0.000	0.000	0.001	0.001
	(0.002)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Observations	2,395,433	2,360,160	2,359,079	2,358,199	2,358,049	2,357,537	2,341,699	2,347,034
R-squared	0.010	0.002	0.002	0.002	0.002	0.002	0.005	0.007
Individual's controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry-specific time trends	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Region-specific time trends	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: DiD estimates comparing treatment and control individuals. Coop equals 1 for individuals employed in a worker cooperative in a particular month and 0 otherwise. The post-reform variable equals 1 for 2011-2013 (policy-on period) and 0 for years 2008-2010 (policy-off period). Individual-level controls include age, male, tenure, firm size (log of total employment), average firm wage (in logs), 9 industry dummies, 19 regional dummies ("Departamentos"). Standard errors clustered at the individual level are reported in parentheses. *** p<0.01, ** p<0.05, * p<0.1

Table 8: Difference-in-differences estimates by firm size

	Incidence of sickness-related absence			Duration (days)		
	(1)	(2)	(3)	(4)	(5)	(6)
	Small firms	Medium firms	Large firms	Small firms	Medium firms	Large firms
Post-Reform × Treatment	-0.002*	-0.002	0.006**	-0.071**	-0.008	0.185**
	(0.001)	(0.002)	(0.003)	(0.033)	(0.057)	(0.074)
Post-Reform × Treatment × Coop	0.009	0.020**	0.034**	0.199	0.445*	0.823**
	(0.007)	(0.010)	(0.015)	(0.177)	(0.259)	(0.403)
Observations	1,184,625	584,913	625,895	1,184,625	584,913	625,895
R-squared	0.006	0.011	0.023	0.005	0.010	0.021
Individual's controls	Yes	Yes	Yes	Yes	Yes	Yes
Industry-specific time trends	Yes	Yes	Yes	Yes	Yes	Yes
Region-specific time trends	Yes	Yes	Yes	Yes	Yes	Yes

Notes: DiD estimates comparing treatment and control individuals. Coop equals 1 for individuals employed in a worker cooperative in a particular month and 0 otherwise. The post-reform variable equals 1 for years 2011-2013 (policy-on period) and 0 for years 2008-2010 (policy-off period). Individual-level controls include age, male, tenure, firm size (log of total employment), 9 industry dummies, 19 regional dummies ("Departamentos"). Standard errors clustered at the individual level are reported in parentheses. *** $p<0.01$, ** $p<0.05$, * $p<0.1$

Table 9: Difference-in-differences estimates: probability of being dismissed

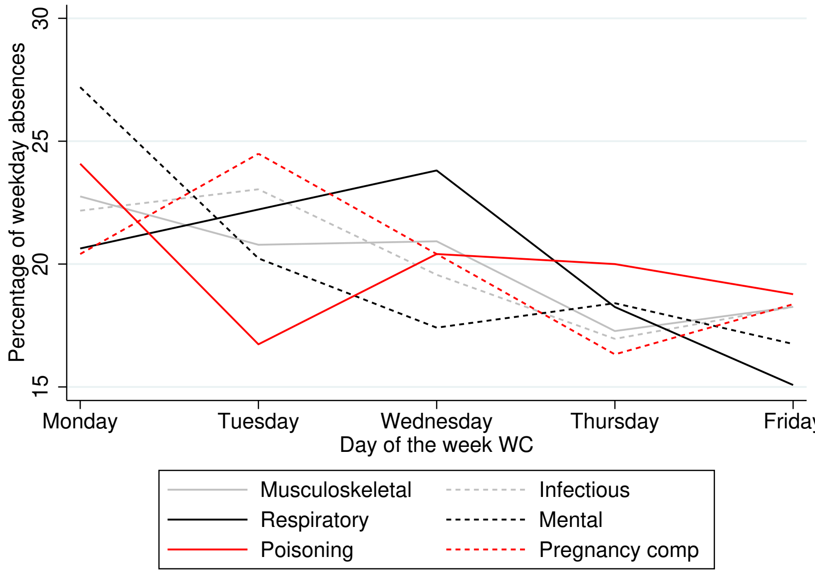
			Small Firms	Large Firms	Only members in worker coops (18-59 years)
	(1)	(2)	(3)	(4)	(5)
Post-Reform × Treatment	0.002***	0.002***	0.001**	0.006***	0.002***
	(0.000)	(0.000)	(0.000)	(0.001)	(0.000)
Post-Reform × Treatment × Coop	-0.007***	-0.008***	-0.002	-0.020***	-0.011***
	(0.001)	(0.001)	(0.001)	(0.004)	(0.002)
Observations	2,362,933	2,362,933	1,169,451	616,779	2,040,093
R-squared	0.005	0.005	0.004	0.010	0.004
Individual's controls	Yes	Yes	Yes	Yes	Yes
Year dummies	Yes	Yes	Yes	Yes	Yes
Industry-specific time trends	No	Yes	Yes	Yes	Yes
Region-specific time trends	No	Yes	Yes	Yes	Yes

Notes: DiD estimates comparing treatment and control individuals. Dependent variable: indicator for whether individual i experienced a layoff in month t. Coop equals 1 for individuals employed in a worker cooperative in a particular month and 0 otherwise. The post-reform variable equals 1 for 2011-2013 (policy-on period) and 0 for years 2008-2010 (policy-off period). Individual-level controls include age, male, tenure, firm size (log of total employment), 9 industry dummies, 19 regional dummies ("Departamentos"). Standard errors clustered at the individual level are reported in parentheses. *** $p<0.01$, ** $p<0.05$, * $p<0.1$

APPENDIX

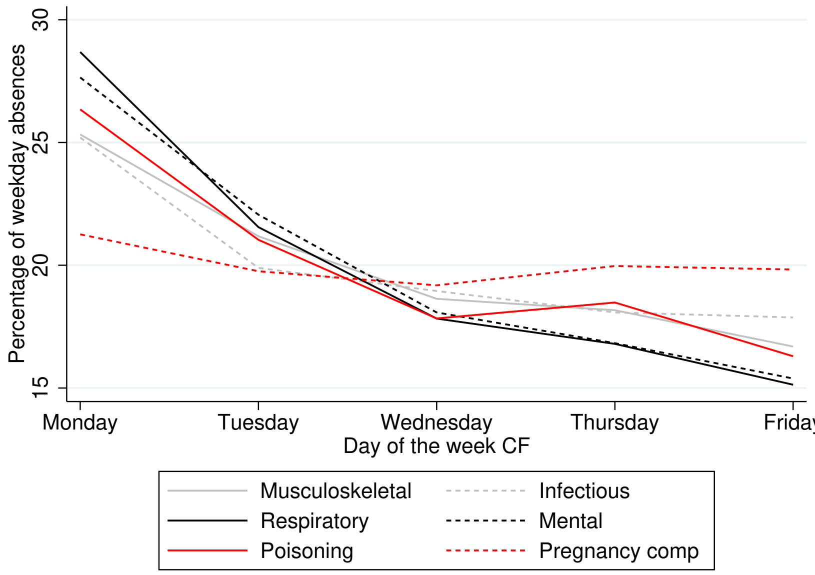
A.1 Additional Results: Figures and Tables

Figure A.1.1: Distribution of sickness-related absence spells by day of first report and disease category (Worker Cooperatives)



Notes: The figure displays the distribution of sickness-absence spells by day of first report and disease category for individuals employed in worker cooperatives (WC).

Figure A.1.2: Distribution of sickness-related absence spells by day of first report and disease category (Conventional firms)



Notes: The figure displays the distribution of sickness-absence spells by day of first report and disease category for individuals employed in conventional firms (CF).

Table A.1.1: Robustness checks: flexible DiD specification

	(1) Incidence	(2) Duration
Post-Reform x Treatment	-0.000	-0.006
	(0.001)	(0.030)
Post-Reform x Treatment x Coop	0.012**	0.302*
	(0.006)	(0.154)
Constant	0.000	-0.115
	(0.015)	(0.203)
Observations	2,395,433	2,395,433
R-squared	0.020	0.016

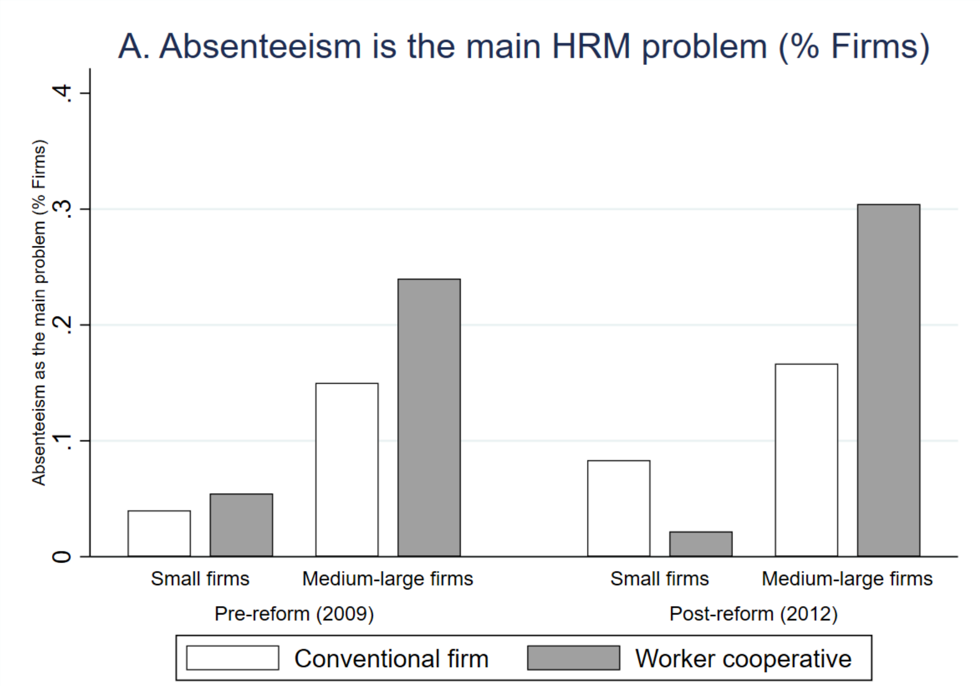
Notes: DiD estimates comparing treatment and control individuals. Coop equals 1 for individuals employed in a worker cooperative in a particular month and 0 otherwise. The post-reform variable equals 1 for years 2011-2013 (policy-on period) and 0 for years 2008-2010 (policy-off period). Individual-level controls include age, male, tenure, firm size (log of total employment), 9 industry dummies, 19 regional dummies ("Departamentos"). Covariates are interacted with treatment status, cooperative status and post-reform period dummies. Standard errors clustered at the individual level are reported in parentheses. *** p<0.01, ** p<0.05, * p<0.1

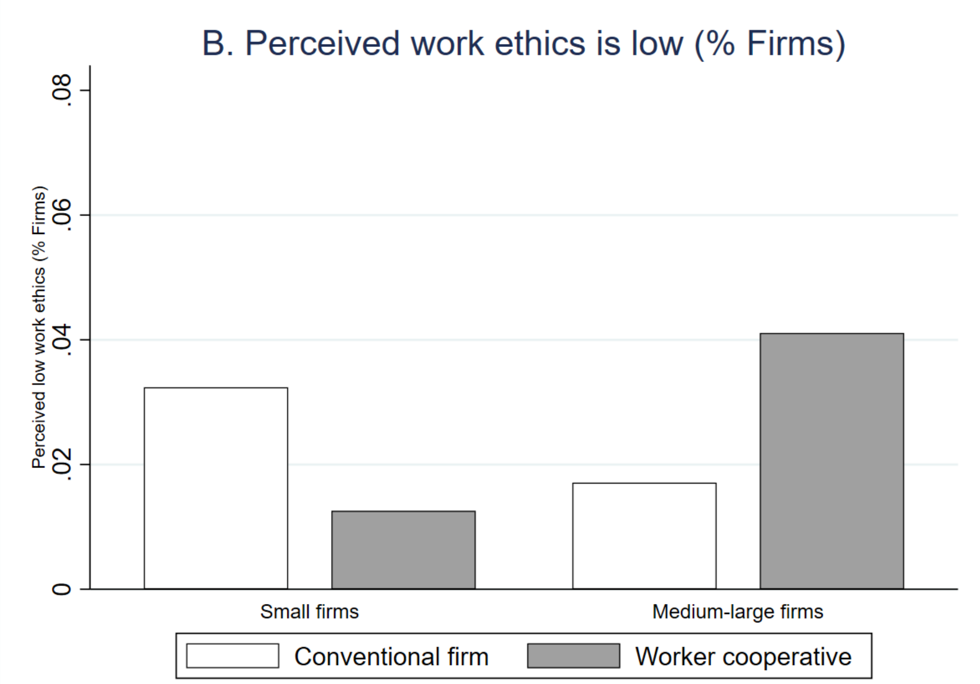
Table A.1.2: Robustness checks: DiD estimates (Controls vs. 3-4 BPCs)

	(1) Incidence	(2) Duration
Post-Reform x Treatment	0.002	0.032
	(0.001)	(0.034)
Post-Reform x Treatment x Coop	0.014*	0.360*
	(0.008)	(0.213)
Observations	1,800,877	1,800,877
R-squared	0.022	0.017
Individual's controls	Yes	Yes
Year dummies	Yes	Yes
Industry-specific time trends	Yes	Yes
Region-specific time trends	Yes	Yes

Notes: DiD estimates comparing treatment and control individuals. Coop equals 1 for individuals employed in a worker cooperative in a particular month and 0 otherwise. The post-reform variable equals 1 for years 2011-2013 (policy-on period) and 0 for years 2008-2010 (policy-off period). Individual-level controls include age, male, tenure, firm size (log of total employment), average firm wage (in logs), 9 industry dummies, 19 regional dummies ("Departamentos"). Treatment group restricted to individuals who were only intended to benefit from the initial sick pay cap hike in January 2011. Standard errors clustered at the individual level are reported in parentheses. *** p<0.01, ** p<0.05, * p<0.1

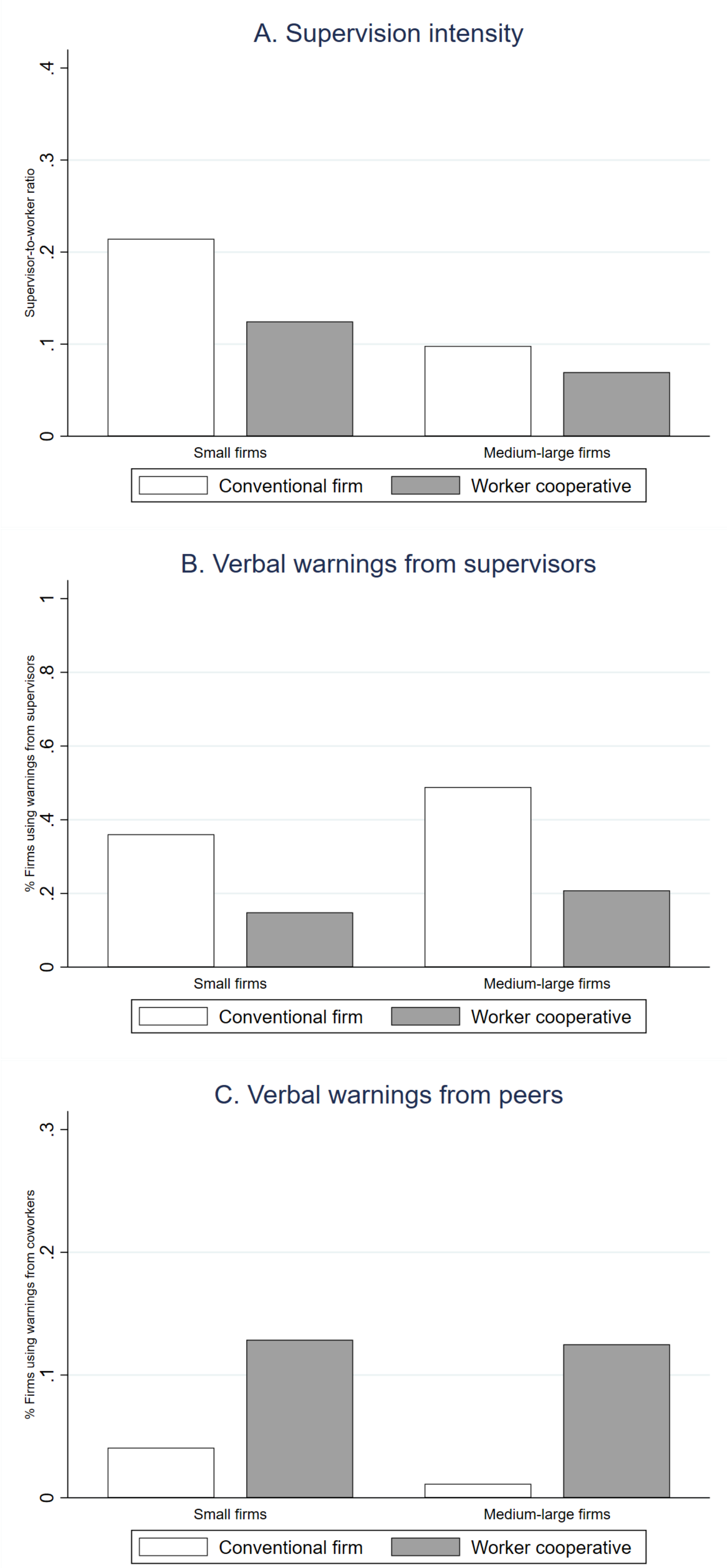
Figure A.1.3: Managers' perceptions about absenteeism by firm size and ownership





Notes: The figure displays the fraction of managers reporting absenteeism as the main HRM problem in the last year (Panel A) and the fraction of managers perceiving that work ethics is low or very low (Panel B). Data from pre-reform (2009) and post-reform (2012) waves of a survey to Uruguayan worker cooperatives and conventional firms of similar size and industry composition. The question on perceived work ethics was introduced in the post-reform wave of the survey. See Section 6 for further details.

Figure A.1.4: Worker supervision and disciplinary mechanisms by firm size and ownership



Notes: The figure displays the supervisor-to-worker ratio (Panel A), the fraction of firms indicating "Verbal warnings from supervisors" as the main disciplinary mechanism (Panel B) and the fraction of firms indicating "Verbal warnings from coworkers" (mutual monitoring) as the main disciplinary mechanism (Panel C). Pooled data from pre-reform (2009) and post-reform (2012) waves of a survey to Uruguayan worker cooperatives and conventional firms of similar size and industry composition. See Section 6 for further details.

Table A.1.3: Heterogeneous effects by age

	Incidence of sickness-related absence			Duration		
	(1)	(2)	(3)	(4)	(5)	(6)
	<35 years	35-49 years	49+ years	<35 years	35-49 years	49+ years
Post-Reform x Treatment	0.003**	0.003	-0.001	0.068*	0.052	-0.006
	(0.001)	(0.002)	(0.003)	(0.035)	(0.055)	(0.068)
Post-Reform x Treatment x Coop	-0.007	0.009	0.028***	-0.139	0.161	0.664**
	(0.008)	(0.012)	(0.011)	(0.177)	(0.329)	(0.284)
Observations	1,147,379	615,668	525,805	1,147,379	615,668	525,805
R-squared	0.020	0.023	0.022	0.016	0.018	0.017
Individual's controls	Yes	Yes	Yes	Yes	Yes	Yes
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes
Industry-specific time trends	Yes	Yes	Yes	Yes	Yes	Yes
Region-specific time trends	Yes	Yes	Yes	Yes	Yes	Yes

Notes: DiD estimates comparing treatment and control individuals. Coop equals 1 for individuals employed in a worker cooperative in a particular month and 0 otherwise.The post-reform variable equals 1 for years 2011-2013 (policy-on period) and 0 for years 2008-2010 (policy-off period). Individual-level controls include male, tenure, firm size (log of total employment), 9 industry dummies, 19 regional dummies ("Departamentos"). Standard errors clustered at the individual level are reported in parentheses. *** p¡0.01, **p¡0.05, * p¡0.1

Table A.1.4: Heterogeneous effects by sector

	Incidence of sickness-related absence				Duration			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Manufacturing	Services	Transport	Others	Manufacturing	Services	Transport	Others
Post-Reform x Treatment	0.011***	-0.000	-0.003	-0.015***	0.250***	-0.009	-0.074	-0.413***
	(0.003)	(0.001)	(0.004)	(0.005)	(0.074)	(0.031)	(0.095)	(0.134)
Post-Reform x Treatment x Coop	0.007	0.015*	0.010	0.048**	0.256	0.278	0.300	1.298**
	(0.014)	(0.008)	(0.015)	(0.021)	(0.393)	(0.216)	(0.408)	(0.594)
Observations	519,070	1,515,733	240,486	120,144	519,070	1,515,733	240,486	120,144
R-squared	0.027	0.017	0.014	0.019	0.022	0.012	0.012	0.017
Individual's controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry-specific time trends	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Region-specific time trends	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: DiD estimates comparing treatment and control individuals. Coop equals 1 for individuals employed in a worker cooperative in a particular month and 0 otherwise. The post-reform variable equals 1 for years 2011-2013 (policy-on period) and 0 for years 2008-2010 (policy-off period). Individual-level controls include age, male, tenure, firm size (log of total employment), 19 regional dummies ("Departamentos"). Standard errors clustered at the individual level are reported in parentheses. *** p¡0.01, **p¡0.05, * p¡0.1

Table A.1.5: Heterogeneous effects by gender

	Incidence of sickness-related absence		Duration	
	(1)	(2)	(3)	(4)
	Male	Female	Male	Female
Post-Reform x Treatment	0.003**	0.005**	0.039	0.112**
	(0.001)	(0.002)	(0.031)	(0.049)
Post-Reform x Treatment x Coop	0.014**	0.011	0.461***	0.129
	(0.006)	(0.011)	(0.170)	(0.296)
Observations	1,289,192	1,106,241	1,289,192	1,106,241
R-squared	0.009	0.023	0.008	0.018
Individual's controls	Yes	Yes	Yes	Yes
Year dummies	Yes	Yes	Yes	Yes
Industry-specific time trends	Yes	Yes	Yes	Yes
Region-specific time trends	Yes	Yes	Yes	Yes

Notes: DiD estimates comparing treatment and control individuals. Coop equals 1 for individuals employed in a worker cooperative in a particular month and 0 otherwise. The post-reform variable equals 1 for years 2011-2013 (policy-on period) and 0 for years 2008-2010 (policy-off period). Individual-level controls include age, tenure, firm size (log of total employment), average firm wage (in logs), 9 industry dummies, 19 regional dummies ("Departamentos"). Standard errors clustered at the individual level are reported in parentheses. *** p¡0.01, **p¡0.05, * p¡0.1

A.1.1 Worker buyouts

As a complementary exercise, we compare absence behaviour before and after the conversion of a conventional firm into a worker cooperative, i.e. a worker buyout. The empirical identification of worker buyouts is not straightforward. For individuals employed in worker cooperatives, we have information about the previous firms at which individuals were employed before joining the worker cooperative.

Following Dean 2019, a worker cooperative that meets the following criteria is considered a worker buyout: (1) more than 50% of the founding members of the worker cooperative were previously employed at the same conventional firm; (2) that conventional firm reduced its workforce by at least 90% either before or in the first operational year of the newly created worker cooperative; (3) both the conventional firm that closed down and the new worker cooperative operate in the same industry. Previous research using similar criteria has identified 58 events of worker buyouts (Dean 2019). For this additional exercise, which is completely independent from our main DiD approach, we only consider 5 worker buyout events that occurred between 2005 and 2013, i.e. the time window for which we have information on sickness-related absences, and restrict the sample to 240 individuals who experienced the two organisational systems at the same firm.

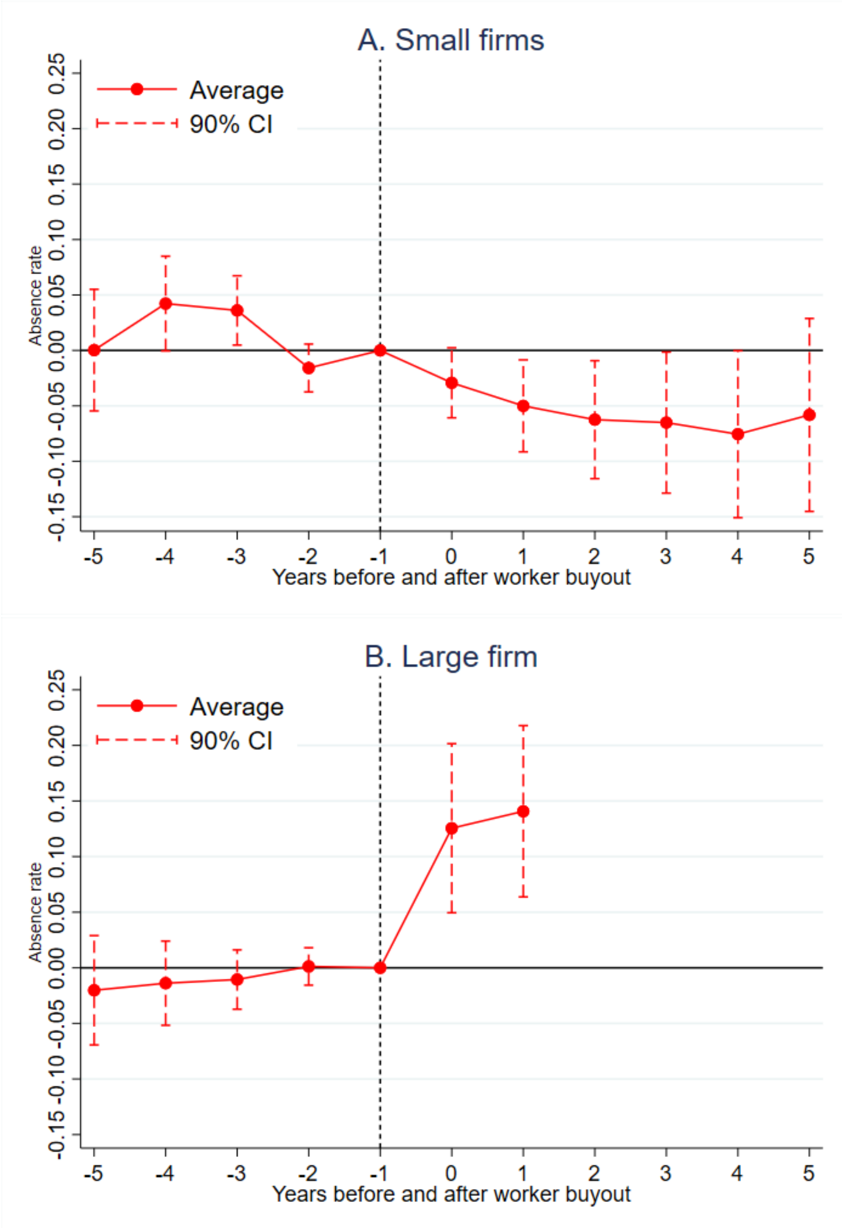
We divide the analysis according to the size of the firms: we observe four worker buyouts of small firms employing less than 20 workers and one worker buyout of a large firm (184 workers). In the latter case, as the buyout occurred in 2011, we can only track individuals under the new cooperative ownership structure for 2 years. To uncover patterns of absence behaviour around worker buyout events, we estimate models of the following form:

$$\mathrm{y}_{it} = \alpha_i + \sum_{j=-5}^{5} \theta_j WBO_{it}^{j} + \beta' X_{it} + \epsilon_{it}$$

where y_{it} either measures whether individual i experienced a sickness absence spell (lasting at least four days) in month t. X_{it} is a vector personal and firm-level characteristics. Our variables of interest are a series of dummy variables WBO_{it}^{j} indicating how many years j it has been since the worker buyout at a given time t. We further include individual fixed-effects α_i to account for time-invariant unobservable characteristics.

Figure A.1.5 displays the estimated coefficients of interest considering the year before the worker buyout as the baseline category. We distinguish the case of worker buyouts of small firms and the worker buyout of a large firm. Interestingly, we observe an asymmetric response of absence behaviour depending on firm size. In the case of the large firm, we find a significant increase in the incidence of sickness-related absences after the buyout (Panel B). By contrast, there is some evidence of a reduction in absenteeism for individuals who experienced a worker buyouts at small firms (Panel A). Although broadly consistent with our main analysis documenting differences between small and large cooperatives, these results should be interpreted cautiously, given the small number of cases.

Figure A.1.5: Incidence of sickness-related absences before and after a worker buyout



Notes: The figures displays the estimated coefficients associated with a vector of dummy variables WBO_{it}^{j} indicating how many years j it has been since the worker buyout at a given time t. See Appendix A.1.1