Switching Costs and Competition in Retirement Investment

By Fernando Luco*

How do different switching costs affect choices and competition in a private pension system? I answer this question in a setting in which variation in employment status allows me to identify two switching costs that jointly affect enrollees' decisions: the cost of evaluating financial information and the cost of the bureaucratic process that enrollees must navigate when switching. I use this variation to estimate the different switching costs and study their impact on competition among pension funds. I find that though eliminating all switching costs decreases equilibrium fees the most, eliminating either switching cost decreases fees significantly.
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In many markets, such as retirement investment and health care, consumers stick to their decisions even when their economic environment changes. Economists often refer to this behavior as "inertia" and explain it as a consequence of switching costs. Though switching costs have been of interest to economists for decades (see Farrell and Klemperer 2007 and the references therein), a number of features of switching costs have often been overlooked. For example, we know little about the nature of the switching costs that affect consumers in actual settings, as these may involve psychological costs, time costs, and penalty fees, among others. In addition, consumers may face several of these switching costs simultaneously, and little is known about their individual effects on competition. It is therefore important to identify the different elements of switching costs and how they affect behavior so as to inform policy design, improve efficiency, and intensify competition.

This paper investigates how inertia affects retirement-investment choices in the Chilean defined-contribution pension system. The Chilean system, which has become a model for many pension systems around the world,[1] provides a key

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[1]These include Argentina, Bolivia, Colombia, Costa Rica, Dominican Republic, Ecuador, Mexico,

advantage for this study because its institutions allow me to separately identify two sources of switching costs that affect enrollees' behavior: the cost associated with analyzing financial information and choosing a pension fund ("decision cost"), and the hassle cost of the time-consuming bureaucratic process that enrollees must navigate to switch pension funds ("enrollment cost"). This allows me to study how policies that eliminate specific switching costs affect enrollees' behavior and competition.

This paper makes three main contributions to the literature. First, I provide evidence regarding the sources of inertia and its costs: Only 44 percent of enrollees switched pension funds between 1988 and 2001, despite significant changes in fees and market structure, and those who switched paid less than those who did not switch, while receiving the same returns.[2]

The second contribution of the paper is to quantify the magnitude of the different switching costs. To do this, I estimate demand for pension funds, taking switching costs into account. The results show that, on average, decision costs are 27 percent larger than enrollment costs, though there is significant heterogeneity across the population.

Finally, the third contribution of the paper is to use counterfactual exercises to study how the different switching costs affect enrollee and firm behavior. The results show that eliminating all switching costs intensifies competition the most and decreases equilibrium fees by 58 percent relative to the case in which enrollees face all switching costs. Eliminating one switching cost at a time also intensifies competition significantly, as eliminating only enrollment costs decreases fees by 41 percent, while eliminating only decision costs decreases fees by 38 percent.

This paper is related to two strands in the literature. The first studies the implications of switching costs for consumer and firm behavior without distinguishing between coexisting switching costs. Goettler and Clay (2011), Nosal (2012), Handel (2013), Polyakova (2016), and Shcherbakov (2016) show that switching costs have a significant impact on consumer choices in contexts such as online retail, health insurance, and paid TV systems, without distinguishing between different switching costs that may affect behavior. Two papers have distinguished between sources other than switching costs that may induce inertia. Miravete and Palacios-Huerta (2014) distinguish between inattention, state dependence, and learning in the context of telephone contracts. Hortaçsu, Madanizadeh and Puller (2017) distinguish between incumbency advantage and consumer inattention in the context of residential electricity markets. Finally, only Dubé, Hitsch and Rossi (2009) has empirically studied how reducing switching costs compares to eliminating them. In the context of two consumption goods, they show that low switching costs induce lower prices than the elimination of switching costs.

Peru, and Uruguay, among others. Also, countries such as Hungary and Poland reformed their retirement-investment systems following some of the ideas on which the Chilean system is based. In addition, the Chilean model has played a role in debates about retirement-investment reforms in many other countries.

[2]That returns do not vary significantly across pension funds is a regulatory feature described in Section I.

The theoretical literature on switching cost and prices, however, has shown that in the presence of switching costs, prices can either increase or decrease (Klemperer, 1987*a*,*b*; Beggs and Klemperer, 1992; Arie and Grieco, 2014; Cabral, 2016).

The second strand of the literature studies investment choices in the context of retirement investment in Chile, Mexico, and the United States. Though this literature has grown significantly in the last years and has documented the existence of inertia, most of it is focused on studying how plan design and information framing affect choices such as participation, asset allocation, and contribution rates (e.g., Madrian and Shea 2001; Hastings and Tejeda-Ashton 2008; Carroll et al. 2009; Choi, Laibson and Madrian 2009; Hastings, Mitchell and Chyn 2010; Hastings 2010; Choi, Laibson and Madrian 2011; Duarte and Hastings 2012; Beshears et al. 2013, 2015; Grubb 2015, among others). This paper contributes to this literature, in that I am able to quantify the magnitude of the switching costs and study their individual impacts on consumer and firm behavior.

Within this strand of the literature, a second line of research has taken a different approach. Hastings, Hortaçsu and Syverson (forthcoming) study investor choices and pricing in the Mexican privatized Social Security System, and show that the interaction between enrollees and the pension fund's sales force is associated with enrollees shifting attention from fees to non-fee fund attributes. Krasnokutskaya, Li and Todd (forthcoming) study how regulation that was meant to limit Chilean enrollees' exposure to risk resulted in pension-fund managers choosing riskier portfolios. Illanes (2016) also studies how switching costs affected enrollees' and firms' behavior, but does so in a later period, when, from a practical perspective, only one of the switching costs studied in this paper remained.[3]

This paper is organized as follows. Section 2 presents a short description of the Chilean retirement-investment system. Section 3 introduces the data and reduced-form evidence of how the different switching costs affect enrollees' behavior and the cost they induce. Section 4 presents the demand model and the identification strategy used to identify the different components of switching costs. Section 5 discusses the main results regarding the magnitude and heterogeneity of switching costs. Section 6 present counterfactuals that study how the different switching costs affect enrollees' decisions and competition among PFAs. Section 7 concludes.

I. The Chilean System

Chile reformed its pension system in several steps, beginning in 1981. The new system covers all Chileans who have worked in the formal sector since 1983.

In the reformed system, enrollees contribute 10 percent of their monthly salary to their private pension account every month and choose a pension fund admin-

[3] A broad body of literature uses Chilean data to estimate demand for pension funds in reduced form (Bernstein and Micco, 2002; Bernstein and Ruiz, 2004; Cerda, 2005; Marinovic and Valdés, 2005; Bernstein and Cabrita, 2006).

istrator (PFA henceforth) to manage their balances.[4]

PFAs can choose the level of fees they charge, but must charge all customers the same fees. Until 2009, these fees were typically a percentage of monthly salary plus a fixed fee. Since 1988, fees can only be charged when enrollees make new contributions and not on existing balances. Figure A.1 in the Online Appendix reports the evolution of quartiles of fees over time, and shows that there has always been considerable dispersion in fees among PFAs.

Between 1988 and 2001, each PFA offered only one investment product, which was composed of different financial instruments. Because Chilean regulation restricts investing to a limited set of financial assets, there is little portfolio differentiation across PFAs.[5] Chilean regulation also explicitly limits enrollees' exposure to risk. In particular, regulation requires PFAs with returns that fall 2 percent or more below the industry mean to cover the losses with their own capital.[6] As a result, PFAs have strong incentives to mimic each other and, as a consequence, there is little dispersion in returns (Figure A.2 in the Online Appendix).

Since 2002, PFAs may offer up to five investment alternatives, and enrollees may divide their balances between no more than two of these alternatives. Enrollees must also remain with the same PFA, who is not allowed to charge differently for the different investment alternatives. Because the data used in this study only allow me to observe which PFA a person is enrolled with, but not in which investment alternative she has invested her funds, I limit the analysis to the period between 1988 and 2001.

Enrollees in the Chilean pension system are susceptible to inertia for at least two reasons. First, they may find it difficult to evaluate the information they have available and choose a new PFA, which creates a tendency to remain with their current PFA. I refer to the cost of evaluating information and making a decision as a "decision cost." Second, to switch PFAs, enrollees must either visit the new PFA or meet with representatives of the new PFA to complete the enrollment process.[7] The cost of the time spent in this enrollment process may also induce enrollees to remain with their current PFA. I refer to this as an "enrollment cost."

To understand why making a decision may be costly, it is necessary to understand the information available to enrollees. Enrollees receive financial information about PFAs and their personal balances in regular reports and personalized statements released by both the regulator of the system and PFAs. In addition, enrollees are exposed to reports, statements, and advertising that provide information about the entire industry along two dimensions: the fees charged by PFAs

[4]The contribution rate is set by law, and has remained at the same level since 1981.

[5]Though investment limits have changed over time, there has always been a limit on fractions of portfolios allocated to risky assets by type and by origin. See Ferreiro (2003), pages 138 and 139, for the evolution of investment limits over time.

[6]Until 2008, if returns were 2 percent or more above the industry mean, the excess returns were saved to cover losses in case the lower bound was reached. This bandwidth is defined by 36-month annualized real returns and is enforced for every month. See Ferreiro (2003), page 82.

[7]Today the enrollment process can be done online. This was not the case during the period studied in this paper.

(a two-part tariff during this period) and measures of real returns for different time periods (year-to-date, last 12 months, last 36 months, last 60 months, and returns since creation of the system).[8] The availability of this type of information suggests that financial literacy may play a significant role during the switching process. Hastings, Mitchell and Chyn (2010) use a randomized experiment to measure Chileans' level of financial literacy and evaluate how they respond to information framing. They show that, similar to what has been found in other countries, financial literacy in Chile is generally low, but increasing in education and income. They also find that people with higher financial literacy choose PFAs based more on past returns and costs than people with lower financial literacy. This suggests that decision costs decrease with financial literacy. On the other hand, because the enrollment process described in the previous paragraph is time-consuming, enrollment costs are likely to be higher among people with higher financial literacy. This suggests that there could be significant heterogeneity in the population regarding the costs that induce inertia.

II. Data and Motivating Evidence

A. Data

I study the period 1988–2001, during which fees could only be charged when enrollees made new contributions and PFAs only offered a single asset allocation. I use an administrative dataset provided by the regulator of the Chilean retirement-investment market, *Superintendencia de Pensiones* (SP). The dataset is a panel of individual histories that starts in 1981 and contains information from the first enrollment until either 2009 or, for older workers, their last contribution. The dataset consists of a random sample of Chileans and was generated in 2002 by the SP. Individual contribution histories were recreated by the SP using PFAs' administrative records.

I use a sample of individuals chosen according to the following criteria. First, because the panel created by the SP does not contain information on the PFA in which an individual is enrolled, but rather only the fees she paid, I use these fees to match individuals with PFAs using auxiliary data regarding the fees each PFA charged in each month during my sample period. As a consequence, I dropped all individuals I was unable to unambiguously match to a PFA. Second, I dropped those individuals who participated in the old pension system because the (voluntary) decision to move from the old system to the new one may have rendered them systematically different from those who entered the system later. Third, I dropped all individuals who entered before 1988 because the fee schedule used between 1981 and 1987 (which allowed PFAs to charge fees based on balances)

[8]The information that PFAs provide to their enrollees in the form of statements and advertising is also regulated to prevent the PFAs from providing misleading information. In general, this means that PFAs must provide the source of the information they report, which is usually a report released by the regulator.

generated a significant number of false matches between individuals and PFAs. Fourth, I dropped the few cases in which the first contribution of an enrollee occurred after the age of 65 or before the age of 15. Finally, I dropped all individuals who had not contributed to their accounts for more than 60 consecutive months. The final sample consists of 350,660 observations of 8,888 people who enrolled between 1988 and 2001, and includes their initial choice and every contribution they made between the moment they entered the system and 2001. In the sample, 635 people on average entered the system each year.

For each person, the dataset includes variables that reflect salaries, fees paid, and characteristics such as gender, birth, enrollment date, and whether the enrollee opened a voluntary savings account for additional savings.[9] I also observe the choice set of each individual, the fees charged by all PFAs, and the returns of the portfolios managed by each PFA. In the analysis, I use average annual real returns computed over 36 months as a measure of past performance.

Table 1 reports summary statistics for three samples that differ on the selection criteria as described in the previous paragraph. The first column reports statistics for people I was unambiguously able to match to PFAs. The second column restricts this sample by dropping people who enrolled before 1988 and also people who, when enrolling, were either younger than 15 or older than 65, both of which are rare. Finally, the estimation sample is obtained after dropping individuals who did not contribute for 60 or more consecutive periods. The table shows that the main impacts of restricting the sample are that in the estimation sample there are a higher percentage of women than in the original sample (50 vs. 44 percent), the mean enrollment age is lower (25 vs. 28), and the fraction of enrollees with voluntary savings accounts is lower (16 vs. 21 percent). Finally, the average monthly salary in the estimation sample is 222,552.4 in 2001 Chilean pesos (332.6 in 2001 U.S. dollars).

Table 1 also shows that in the Chilean system, people can go long periods without contributing to their accounts, even though the mean number of months between the first and last recorded contributions is high and participation is mandatory for formal employees. This phenomenon is well known, and could be caused either by long unemployment spells, frequent transitions between employment and unemployment, transitions between formal and informal employment, long periods of informal employment, or transitions between formal employment and self-employment. The data do not allow me to distinguish between these cases. However, as individuals in other employment categories rarely participated in the system during this period, the assumption that all observed contributions correspond to formally employed individuals is not strong.[10] Finally, I use the

[9] In addition to saving for retirement using their mandatory account, enrollees can save additional resources using a separate voluntary savings account. The voluntary account is managed by the same PFA as the mandatory account and invested in the same instruments. The data available only identify whether the enrollee opened a voluntary account but not the moment at which the account was opened.

[10] Independent workers between 1988 and 2002 accounted for 2.9 percent of total enrollees. See the series “Afiliados por tipo y sexo” (Enrollees by type and gender) in

TABLE 1—: Summary statistics by sample

	Sample 1	Sample 2	Estimation Sample
Number of people	15,458	9,257	8,888
Gender (Female)	0.44	0.5	0.49
Age when enrolling	27.98	24.65	24.66
	(10.26)	(8.89)	(8.92)
Age in 2001	36.35	30.74	30.58
	(12.02)	(9.7)	(9.71)
Percentage of enrollees with voluntary savings	0.21	0.16	0.16
Number of months observed	89.57	66.12	63.94
	(58.11)	(48.68)	(48.17)
Income in December 2001 pesos	247,073.9	221,901.1	222,552.4
	(490,296.8)	(238,748.1)	(239,173.8)
Time inactive (months)	15.77	14.61	11.85
	(21.01)	(19.49)	(13.82)
Number of changes	30,091	12,560	12,196
(percentage of total decisions)	3.4	3.5	3.5
Number of observations	886,087	358,800	350,660

Note: The table reports means and standard deviations (in parentheses) for selected demographics across the different samples that result from applying the selection criteria described in Section I. Sample 1 corresponds to people who are unambiguously matched to a PFA. Sample 2 drops people who enrolled before 1988 and those who, when enrolling, were younger than 15 or older than 65. The estimation sample drops people who did not contribute to their accounts for more than 60 consecutive months.

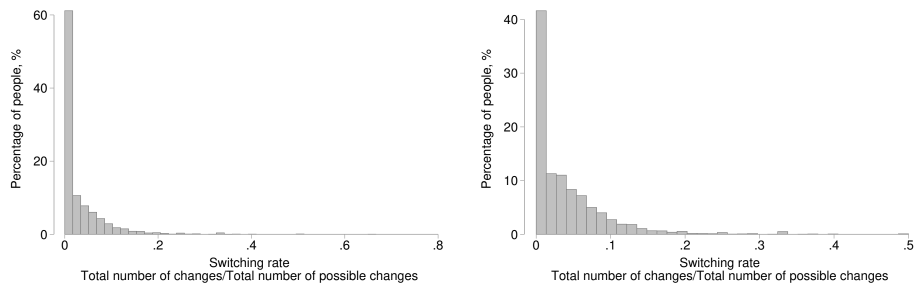
first contribution after a period with no contributions to identify when an enrollee returns to the system.

B. Motivating Evidence: Inertia, Its Cost, and Who Switches

I begin by showing that enrollees bear a cost associated with inertia in the form of higher fees paid. Second, having shown the costliness of inertia, I investigate when people switch and the attributes of switchers. The results show that people who return to the system after periods in which they did not contribute are significantly more likely to switch PFAs than enrollees who have contributed to their accounts continuously, even after controlling for numerous demographics.

Before turning to the evidence, it is important to know whether Chileans switch pension funds at all. Figure 1 reports the switching rate–the ratio between the number of times an individual switched pension funds and the number of times this individual made a contribution. The figure shows that switching is rare, though it is more common among those who have been enrolled longer.

http://www.spensiones.cl/safpstats/stats/.sc.php?_cid=43 for more details.



(a) Whole sample (b) Enrollees who entered on or before 1995

Figure 1. : Switching rate.

Note: The figure on the left reports the switching rate for the whole sample, and the figure on the right for those who enrolled for the first time on or before 1995.

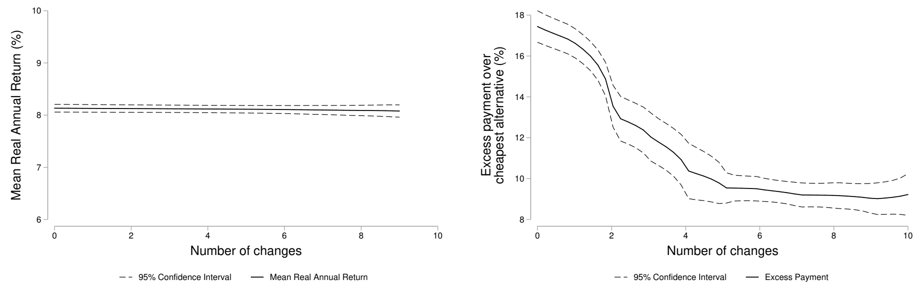
The Cost of Inertia. — To determine whether inertia has a cost, I focus on how returns and fees vary with the number of times people switch. Figure 2a reports a local polynomial regression of realized returns by the number of times enrollees switched PFAs. The figure shows that returns do not vary with the number of times enrollees switch. This is a consequence of regulation, and it suggests that enrollees should minimize the fees they pay. Figure 2b shows that though this is not the case, enrollees who switch more tend to choose cheaper options.[11] The figure reports fees paid in excess of the cheapest fund, and shows a negative relationship between fees paid and the number of times people switch.[12] This means that while those who switched five or more times paid 9 percent more than with the cheapest fund, those who never switched paid, on average, 18 percent more than with the cheapest fund.

When Do People Switch, and Who Switches?. — As described above, there are two possible reasons behind the low switching rate. First, enrollees may find it difficult or costly to choose a pension fund because they must analyze financial information in a complex environment. Second, when switching, enrollees must go through a time-consuming enrollment process. These costs may induce enrollees to stay with their current PFA.

An important characteristic of this setting, to which I will return in Section III when discussing the identification of switching costs, is that various groups of enrollees face different switching costs. Enrollees who enter the system for the first time do not have a default option; they must choose a PFA and go through the

[11] Enrollees may also choose funds because of non-fee fund attributes such as returns, their perception of a funds' sales force and customer service, and the length of time a fund has been active, among others. This is considered in estimation. This section, however, focuses on realized returns and fees paid only.

[12] Because participation is mandatory, I measure the cost of inertia relative to the cost of choosing the cheapest fund available.



(a) Excess returns (b) Excess payment

FIGURE 2. : Returns and overpayment by number of changes in PFAs

Note: The figures report local polynomial smoothing of payments in excess of the cheapest option and realized returns as a function of the number of times that enrollees switch PFAs.

enrollment process. Enrollees who return to the system after an absence, called "returning enrollees," must go through the enrollment process even if they do not wish to switch PFAs. This is because when enrollees return to the system—after, for example, being unemployed—they must provide evidence of prior enrollment, even if they decide to remain with their current PFA. Hence, upon returning to formal employment or when switching jobs, the enrollment process is unavoidable. In contrast, enrollees who have contributed continuously, called "existing enrollees," only face enrollment costs if they switch PFAs. Therefore, if the enrollment cost is large enough, it may induce existing enrollees to remain with their current PFA, but it should not affect switching behavior for returning enrollees. This suggests that switching should be more common among returning enrollees than among existing enrollees. I study this hypothesis in this section.

Table 2 shows that 10.9 percent of observations from returning enrollees involve switching PFAs, compared to only 2.5 percent of observations from existing enrollees. Restricting the sample to enrollees who own a voluntary savings account results in similar findings.

Though the findings described in the previous paragraph are compelling, other factors could also affect switching, which could lead to overestimating the effect of returning to the system. Examples include changes in salary when an enrollee returns to the system, or getting closer to retirement. To take this into account, I report the estimates of several specifications of probit regressions in Table 3, in which I study how the probability of switching is affected by different sets of regressors.

The first specification includes an indicator that is equal to one if the enrollee returned to the system in the month of that observation. The effect of returning is both positive and highly significant. The estimated coefficient translates into an 8.44 percentage point higher probability that a returning enrollee will switch

Table 2—: Fraction of observations that involve switching behavior by contribution status

Contribution status	Number of observations	Number of PFA switches	Percentage of switches All enrollees	Percentage of switches Enrollees with Savings Account
Returning enrollees	43,648	4,772	10.9	15.0
Existing enrollees	298,124	7,424	2.5	3.3

Note: The table presents the number of observations, number of changes in PFA, and the number of changes as percentages of observations for existing and returning enrollees. The table was constructed using all observations except initial choices.

than an existing enrollee will switch.[13] The second column replaces the indicator for returning with an indicator that is equal to one if the enrollee had an increase in salary of 10 percent or more relative to her average income over the last five months during which she was an existing enrollee. The result, positive and significant, disappears once we add the indicator for a returning enrollee (column 3). Column 4 shows that the results remain unchanged when year fixed effects are included. Column 5 adds demographic information, and shows that older people switch less frequently, that people with a voluntary savings account switch more often than those without one, and that gender does not affect the probability of switching.[14] The marginal effect is almost identical to that of the first specification. Finally, column 6 shows that the results remain unchanged when PFA fixed effects are included. In summary, the results show that enrollees returning to the system are more likely to switch PFAs than enrollees who have contributed continuously, consistent with these enrollees' decisions not being affected by enrollment costs.

The specifications presented in Table 3 investigated whether returning to the system had an impact on the probability of switching in the same month. To determine whether the influence of returning to the system is limited to the first month after returning or lasts longer, I replicate the last specification in Table 3, but replace the indicator for returning with a variable that measures the number of months since an individual returned to the system. Figure 3 shows that the most significant impact on the probability of switching occurs when enrollees return to the system and the impact vanishes after the second month.

[13]In what follows, all marginal effects are significant at the 1-percent level.

[14]This last result, which is somewhat surprising, as it is known that women are less active than men in the Chilean pension system, can be explained by two factors. First, the sample of enrollees whose decisions are studied in this paper consists of people that have not gone through 60 or more consecutive months without contributing to their accounts. This eliminates enrollees who may have left the system and have inactive accounts. Second, the analysis presented here refers to enrollees' choices that are conditional on contributing. Hence, the women in the sample whose decisions we study are likely to be more active than those who are left out of the sample.

TABLE 3—: Effect of changes in monthly salary and demographics on the probability of switching

	(1)	(2)	(3)	(4)	(5)	(6)
Returning	0.732		0.729	0.736	0.744	0.784
	(0.01)		(0.01)	(0.011)	(0.011)	(0.011)
Increase in salary greater than 10 percent		0.123	0.013	-0.006	-0.015	-0.008
		(0.009)	(0.009)	(0.01)	(0.01)	(0.01)
Age					-0.007	-0.009
					(0.001)	(0.001)
Male					0.003	0.015
					(0.014)	(0.013)
Has a voluntary savings account					0.172	0.151
					(0.015)	(0.014)
Marginal effect of returning (pp)	8.44	-	8.04	8.01	8.08	8.19
Year fixed effects	No	No	No	Yes	Yes	Yes
PFA fixed effects	No	No	No	No	No	Yes
N	341,772	341,772	341,772	341,772	341,772	341,769
Pseudo R^2	0.054	0.002	0.054	0.0963	0.102	0.149
Log Likelihood	-49,809.5	-52,530.4	-49,808.5	-47,557.7	-47,283.3	-44,755.2
χ^2	5,339.9	194.8	5,454.1	6,578.3	6,774.3	11,776.5

Note: Standard errors, clustered at the individual level, in parentheses. An observation is an individual–month combination. The dependent variable is equal to one if the individual switched and zero otherwise. Estimation is by maximum likelihood. The specified model is a probit model. Marginal effects refer to the increase in the probability of switching, in percentage points, associated with the returning indicator taking the value of one relative to zero. All marginal effects are significant at the 1 percent level.

Overall, the evidence shows that returning enrollees are significantly more likely to switch PFAs than existing enrollees, consistent with enrollment costs not affecting decisions of returning enrollees. In Table B.1 in the Online Appendix, I study the robustness of the results presented in this section to the inclusion of additional regressors in the analysis. The results remain unchanged.

III. Model, Identification, and Supporting Evidence

A. *Model*

Consider a model of consumer behavior in which enrollees' decisions are affected by switching costs that depend on employment status. Enrollees must decide each

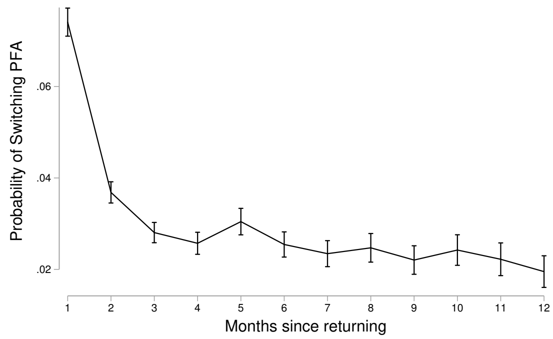


FIGURE 3. : Marginal effect on the probability of switching by month since returning

Note: The figure reports the probability of switching PFA (and the 95 percent confidence interval) as a function of the number of months since an enrollee returned to the system. The specification replicates that in column 6 of Table 3, but replaces the indicator of returning by a variable that measures the number of months since an individual returned.

period whether to remain with their current PFA or switch. Also assume that enrollees are myopic, as will be discussed below. I model utility using a traditional discrete choice approach in which consumers have heterogeneous valuation for product attributes. In particular, I consider fees and returns to be product characteristics, and these characteristics are perceived by consumers through the information that is publicly available about the funds.

Enrollees' behavior is driven by the information they have regarding the PFAs and their individual characteristics. Let y_{it} denote enrollee's i salary, f_{jt} the fixed fee, and p_{jt} the percentage fee charged by pension fund j in period t, and let R_{jt} be an observable measure of returns for fund j in the same period. Let η_{it} denote the cost of switching to fund j in period t. η_{it} represents the sum of the two components, enrollment costs η_{it}^E and decision costs η_{it}^D, depending on the enrollee's employment status. That is, $\eta_{it} = \eta_{it}^D + \mathbf{1}[i \text{ is an existing enrollee at time } t]\eta_{it}^E$. Finally, because enrollees' choices may also be affected by other characteristics of the pension funds, I denote pension fund j's unobserved characteristics as ξ_j (unobserved to the econometrician) and ε_{ijt} to be an idiosyncratic taste shock to consumer i at time t. I assume this shock is identically and independently distributed as an extreme value type I. I discuss these assumptions below.

Under these assumptions, the indirect utility of consumer i when choosing alternative j in period t is given by

$$u_{ijt} = \alpha_{it}(y_{it} - 0.1y_{it} - f_{jt} - p_{jt}\min\{y_{it}, \bar{y}_t\}) + R'_{jt}\beta_{it} - \eta_{it}\mathbf{1}[d_{it-1} \neq j] + \xi_j + \varepsilon_{ijt}. \quad (1)$$

In this setting, consumer i chooses fund j if $u_{ijt} \geq u_{ikt}$ for all $k \neq j$.[15]

[15] I do not include any measure of risk as an observable characteristic of the funds, for three reasons.

The term $y_{it} - 0.1y_{it} - f_{jt} - p_{jt}\min\{y_{it}, \bar{y}_t\}$ represents disposable income after making a contribution to the enrollee's pension account and net of fees. In this expression, fees correspond to the sum of the fixed fee f_{jt} and the percentage fee charged over income p_{jt}, which is multiplied by a function that defines an upper bound on how much enrollees can pay ($\min\{y_{it}, \bar{y}_t\}$). The upper bound $\bar{y}$ is set by law and adjusted by inflation.[16] In this context, α_{it} represents the marginal utility of disposable income.

The second term in Equation 1 refers to how returns affect utility and decisions. While returns could well affect choices, Cerda (2005) argues that due to the small dispersion in realized returns, enrollees in the Chilean system react more to changes in the *ranking* of realized returns than to changes in the returns themselves. PFAs also commonly focus on their return ranking in their advertising. Accordingly, I follow Cerda (2005), and in the main specification I use the PFA's return ranking position to measure how realized returns affect choices. In this context, β_{it} measures how higher positions in the ranking of returns affect utility for consumer i at time t. In Section IV, I study the robustness of the results by using alternative measure of returns, which do not materially change the results.

Though the proposed model is simple and it allows for estimating the distributions of preferences and switching costs, some assumptions must be discussed. First, I assume that enrollees are myopic. This implies that the utility function is defined in reduced form, because enrollees derive flow utility from returns of their pension accounts, even though these payoffs will only be realized upon retiring. This, however, is based on the notion that the best prediction enrollees can make regarding future fees and returns is precisely the one that they currently observe. Furthermore, I do not consider that enrollees may anticipate either changes in fees or the direction of these changes. If this is the case, switching costs would be overestimated relative to flow utility, because while enrollees may wait for these changes to be realized, the model would rationalize this as inertia caused by switching costs.

Although these are limitations of a static model, I believe the static framework represents the Chilean retirement-investment system better than a dynamic framework for two reasons. First, market structure and fees changed significantly and often during the period of analysis. Indeed, the number of PFAs went from

First, as explained earlier, because of the regulatory framework there is little dispersion and no persistence in returns, which essentially eliminates any persistent risk associated with a specific PFA. Second, and more importantly, enrollees do not have easy access to information that might allow them to infer any measure of risk. Indeed, enrollees receive information about the annualized returns for different time periods and the fees that each PFA charges. To compute measures of volatility, an enrollee would need to request the time series of returns from the regulator and compute volatility herself (or collect all of the statements she has received over a time period, extract the information on returns, and compute, for example, the variance). Finally, Hastings, Mitchell and Chyn (2010) show that financial literacy is low among enrollees in the Chilean system, which makes it unlikely that they would know how to compute measures of risk.

[16] In the computation of fees, income is bounded above by a value that changes monthly. During this period, the reference income was defined as 60 U.F. (almost \$1,500 U.S. dollars as of December 2001), where U.F. is a monetary unit of constant value (adjusted by inflation).

12 in 1988 to 22 in 1994 to 7 in 2001, making it difficult for enrollees to form expectations of how the industry, and the fee schedule, would look in the future. Second, regulation concerning both the financial instruments that PFAs could use to construct their portfolios and the investment limits they had to follow also changed significantly and often during this period. These changes introduced complexities that make it unlikely that consumers would follow a dynamic approach when choosing PFAs. It is important to note, however, that this is a different setting from that studied by Illanes (2016), as he focuses on a later period–in which market structure was significantly more stable–which renders a dynamic framework better suited than it would be for earlier periods. Nonetheless, to allow for changes in preferences over the life cycle, the random coefficients vary with demographics such as age and income.[17]

Second, I have assumed that unobservable fund characteristics are constant over time. This assumption, though strong, is necessary for two reasons. First, it is not feasible to allow for time-varying fund unobservables using period-fund fixed effects. Second, an alternative procedure would be to use the "BLP inversion" (see Berry, Levinsohn and Pakes, 1995). However, this requires that all PFAs have positive market shares in all periods, which does not hold in the data, particularly in the first few months after a new PFA enters the market. For this reason, instead of dropping these observations from the data, I estimate ξ_j as fixed over time, but interact it with individual demographics (that do change over time) to capture heterogeneity in brand preferences.

B. *Identification*

The main identification challenge of this paper, as found throughout the empirical literature on switching costs, is to distinguish unobserved preferences from switching costs, as their impact on choice behavior is observationally equivalent. To identify the different switching costs, I use the identification strategy employed by Handel (2013), which leverages the existence of different groups of enrollees, and extend it to exploit the existence of three of these groups.

Specifically, in the Chilean pension system, enrollees face different switching costs depending on their employment status. Enrollees who have continuously contributed to their accounts can avoid decision and enrollment costs by remaining enrolled in their current PFA. On the other hand, enrollees who return to the system after a period of absence incur enrollment costs regardless of whether they decide to remain with their current PFA because they must go through the enrollment process again. Hence, enrollment cost should not affect their decisions. Nonetheless, if decision costs are large enough, they may induce returning enrollees to remain enrolled with their current PFA. This means that enrollment costs can be identified by comparing the choices of existing and returning en-

[17] Another (practical) benefit of the static model is that it allows me to study dynamic price competition in the presence of switching costs in a richer way than possible with a dynamic-demand framework.

rollees.

To identify decision costs, I exploit the existence of a third group: "new enrollees." New enrollees are those who are entering the system for the first time and do not have a default option.[18] For new enrollees, both costs are unavoidable, as they have to choose and enroll with a PFA. This means that neither cost can affect their choices, because they are incurred regardless of the alternative chosen. Therefore, decision costs can be identified by comparing the choices of new and returning enrollees.

Finally, the identification of taste coefficients follows standard arguments. In particular, taste coefficients are identified using initial choices–when all costs are unavoidable and should not affect enrollees' decisions–and exploiting variation in choice sets, prices, and characteristics. The remaining choices identify the different components of switching costs as described above.

C. Supporting Evidence

For the identification strategy to be valid, new, returning, and existing enrollees must have the same preferences, so that differences in choice behavior can only be explained by differences in the switching costs that enrollees face and not by persistent differences across people in different groups. To rule this out, I offer two arguments that suggest that what drives the differences in behavior across groups is switching costs rather than unobserved heterogeneity. In addition, Table C.1 in the Online Appendix compares the demographic composition of the different groups and shows that, with two exceptions, the groups are remarkably similar.[19]

First, almost everyone in my sample was a returning enrollee at least once (82 percent of the people in the sample), and among those who were never returning enrollees, more than half entered the system in the last two years of my data (Figure A.3 in the Online Appendix). The mean number of times an enrollee returned to the system is 5, and the median 4. It is therefore unlikely that returning and existing enrollees consistently differ in their preferences, since enrollees frequently transition between the status of returning enrollee and existing enrollee.

Second, I present reduced-form evidence that compares choice behavior among observationally equivalent enrollees who left the market at the same time but returned in consecutive periods facing different fees. To do this, I estimate probit regressions similar to those reported in Table 3, but restrict the sample to indi-

[18]Until 2009, new enrollees had to chose a PFA when entering the labor market for the first time. This changed in 2009, when as part of an important reform to the system, a mandatory default PFA for new enrollees was introduced.

[19]The exceptions correspond to gender composition and whether the enrollee has a voluntary savings account. Indeed, while 50 percent of the people in the estimation sample are females, males are overrepresented in the returning group. This is associated to specifics of the labor market, however, rather than to selection in the pension-funds market. Finally, returning enrollees are more likely to have a voluntary savings account than existing enrollees. This suggests that there might be further differences across individuals that must to be taken into account. I do so in estimation by controlling for observable individual characteristics–such as age, income, gender, and whether an enrollee has a voluntary savings account–and also by incorporating time-invariant individual-specific unobserved heterogeneity.

viduals who return in consecutive periods ($t-1$ and t), when the fee schedule changes in t relative to that of $t-1$. The idea behind this test is that if switching is caused by changes in preferences rather than by switching costs, someone who returns to the system in $t-1$ should be equally likely to switch in period t, as the individual who returns in t and faces a new fee schedule. On the other hand, if enrollment costs do not affect decisions when returning to the system, those who return in t should be more likely to switch than those who returned in $t-1$.[20]

The results are reported in Table 4. Different columns restrict the sample to periods in which only the fixed fee changed (columns 1 and 3) or the percentage fee changed (columns 2 and 4). Columns 3 and 4 also include fixed effects for the time when an enrollee left the system, to take into account that people who left the market at different times did so under different economic conditions, which could affect future labor-market outcomes. Across columns, the results show that similar enrollees who returned to the system in consecutive periods behave differently when facing the same change in relative fees. This suggest that inertia is not determined by differences in preferences, but rather by the different switching costs faced by each group.[21]

IV. Specification, Estimation, and Results

A. *Specification and Estimation*

As described in the previous section, I specify taste coefficients and switching costs as linear functions of observable demographics and enrollee-specific time-invariant unobservables. Specifically, let D^{α}_{it}, D^{β}_{it}, and D^{η}_{it} be vectors of demographic information for enrollee i in period t, which are used to model both tastes and switching costs. In the baseline specification, I assume that

$$
\begin{aligned}
D^{\alpha}_{it} &= [\text{age}_{it}, \text{gender}_i] \\
D^{\beta}_{it} &= [\text{age}_{it}, \text{gender}_i, \text{income}_{it}] \\
D^{\eta}_{it} &= [\text{age}_{it}, \text{gender}_i, \text{income}_{it}, \text{voluntary savings}_i, \text{regulation}_t].
\end{aligned}
$$

Taste coefficients and switching costs are then parameterized as

[20] Ideally, one would perform this analysis by restricting the sample to enrollees who, when returning in $t-1$, faced the same fees as when they left the system. Therefore, in t both those who returned in $t-1$ and those who returned in t would face the same change in relative fees. In practice, however, this is not feasible due to data limitations. Instead, I restrict the sample to individuals who did not switch PFAs in $t-1$, so that after the change in fees that occurs in t, they face the same change in relative fees as those returning in t.

[21] A caveat regarding this analysis is that it is possible there is selection into when to return to the market. In other words, the analysis would be undermined if enrollees decided to return in t instead of $t-1$ because of the fees they would face in t relative to those of $t-1$, as among other factors, the PFAs may change fees based on general economic conditions that may also influence when an enrollee returns to the system.

TABLE 4—: Effect of returning on the probability that returning enrollees switch when fees change

	Fee changing in month t relative to $t-1$			
	Fixed	Percentage	Fixed	Percentage
	(1)	(2)	(3)	(4)
Returning in t	0.672	0.514	0.887	0.648
	(0.144)	(0.0854)	(0.335)	(0.166)
Age	-0.0112	-0.00415	-0.0146	-0.00581
	(0.00612)	(0.00483)	(0.00839)	(0.00517)
Male	-0.0271	-0.00375	-0.157	-0.0277
	(0.117)	(0.0773)	(0.138)	(0.0876)
Has a savings account	0.164	0.123	0.113	0.0691
	(0.136)	(0.0882)	(0.163)	(0.0998)
Salary	0.00464	0.00392	0.00516	0.00459
(tens of thousands)	(0.00164)	(0.00108)	(0.00186)	(0.00118)
Account balance	0.0000674	-0.0000682	0.000449	0.000392
(tens of thousands)	(0.000424)	(0.000232)	(0.000405)	(0.000211)
Marginal effect (percentage points)	6.15	5.02	9.70	5.76
N	1363	3309	1003	2720
Pseudo R^2	0.0765	0.0469	0.2231	0.1955
Log likelihood	-283.71	-660.82	-207.12	-522.39
χ^2	45.29	58.57	127.34	258.51

Note: Standard errors, clustered at the individual level, in parentheses. The dependent variable is an indicator that is equal to one if the enrollee switches managers in that period and zero otherwise. A unit of observation is an enrollee in a month. Estimation corresponds to a probit regression. The sample corresponds to individuals who returned in consecutive periods ($t-1$ and t), when fees change in t relative to $t-1$. The sample is restricted to individuals who, if returning in $t-1$, did not switch PFAs at that time, and compares their behavior in t with that of those who returned in period t. Columns 3 and 4 repeat regressions presented in columns 1 and 2, adding fixed effects for the time when the enrollee left the market before she returned. Marginal effects refer to the increase in the probability of switching, in percentage points, associated with the returning indicator taking the value of one relative to zero. All marginal effects are significant at the 1 percent level.

$$
\begin{aligned}
\alpha_{it} &= \alpha_0 + D_{it}^{\alpha}\boldsymbol{\alpha} + \sigma_\alpha \mu_i^\alpha \\
\beta_{it} &= \beta_0 + D_{it}^{\beta}\boldsymbol{\beta} + \sigma_\beta \mu_i^\beta \\
\eta_{it}^k &= \eta_0 + D_{it}^{\eta}\boldsymbol{\eta} + \sigma_\eta \mu_i^{\eta_k}; \quad k = \{D, E\} \\
\mu_i^l &\sim N(0,1), \quad l = \{\alpha, \beta, \eta^D, \eta^M\},
\end{aligned}
$$

where D_{it}^{α}, D_{it}^{β}, and D_{it}^{η} are the vectors of demographics introduced above and μ_i^l represents individual-specific time-invariant unobservable characteristics. I use Halton draws to simulate all μ_i^l's and numerically integrate over them.[22]

Because the data cover 14 years, I allow variables such as age and income to change over time. This means that switching costs (η_{it}) and taste coefficients (α_{it} and β_{it}) change over time as well. I specify different vectors of demographics for the taste coefficients, because if income were to be included in the parameterization of α, it is possible that if both income and fees change at the same time, disposable income (income after paying fees) might not change, but α would change mechanically. On the other hand, one could specify α as a function of disposable income rather than income before paying fees, but this would result in $\alpha_{i\cdot}$ varying across options in the choice set. For these reasons, I do not include income in the specification of α.

Under the assumption that taste shocks are *i.i.d* extreme value type I, the probability that enrollee i chooses alternative j in period t is given by

$$\text{(2)} \quad \boldsymbol{P}_{ijt} = \int \frac{\exp(-\alpha_{it}(f_{jt} + p_{jt}\min\{y_{it}, \bar{y}_t\}) + R'_{jt}\beta_{it} - \eta_{it}\mathbf{1}[d_{it-1} \neq j] + \xi_j)}{\sum_k \exp(-\alpha_{it}(f_{kt} + p_{kt}\min\{y_{it}, \bar{y}_t\}) + R'_{kt}\beta_{it} - \eta_{it}\mathbf{1}[d_{it-1} \neq k] + \xi_k)} dF_{\mu},$$

where F_{μ} represents the joint distribution of μ_i's.

Finally, a normalization is required because there is no outside option. I assume that ξ_k is equal to zero for one of the PFAs that was always present in the market. With this, the simulated maximum likelihood estimator $\hat{\theta} = \{\hat{\eta}, \hat{\alpha}, \hat{\beta}, \hat{\xi}, \hat{\sigma}\}$ is given by

$$\text{(3)} \quad \hat{\theta} = \arg\max_{\theta \in \Theta} \sum_i \sum_k \sum_t \log(\boldsymbol{P}_{ikt})\mathbf{1}[d_{it} = k].$$

B. Main Results

Table 5 reports the estimated coefficients of the distributions of switching costs for the main three specifications. The first column does not include unobserved heterogeneity, and the last two do. The difference between columns 2 and 3 is that column 3 includes the interaction between brand fixed effects and individual demographics to account for heterogeneity in brand preferences. I will discuss the results based on this third specification.

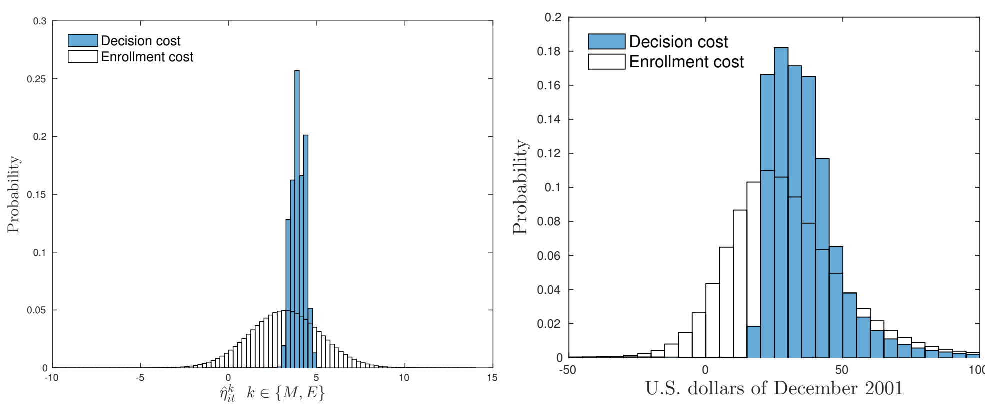
As the table shows, the results are similar across specifications. For both deci-

[22] D_{it}^{η} includes a dummy variable denoted "regulation" that identifies the period 1998–2001. In 1997, the Chilean government passed a regulation that made it harder for enrollees to switch, in response to a number of reports that showed that sales people would offer gifts to enrollees who agreed to switch PFAs. The implemented regulation increased the bureaucratic hassle cost associated with the switching process. The 'Regulation' dummy allows me to take this change into account.

sion and enrollment costs, the constant captures the mean of the distribution of switching costs and the demographics rationalize heterogeneity around the mean. Both switching costs increase with age, are not affected by gender, and decrease with income. In addition, people with voluntary savings accounts have lower decision costs–but higher enrollment costs–than those without a voluntary savings account. This finding is consistent with higher-income enrollees, who are more likely to have a voluntary savings account, being more financially literate and having a higher opportunity cost of time. Finally, the 1997 regulation that made switching more difficult, is associated with higher enrollment and decision costs.

Estimated values of σ are significant for both switching costs. However, σ_E is an order of magnitude larger than σ_D, suggesting that demographics cannot explain as much heterogeneity in enrollment costs as they do for decision costs. Finally, the introduction of unobserved heterogeneity has little impact on the estimates of decision costs. This is not, however, the case with enrollment costs. Indeed, the constant of enrollment costs more than doubles relative to the specifications that do not consider unobserved heterogeneity.

Figure 4 plots the distribution of switching costs by component for the preferred specification (column 3 in Table 5). The figure shows that decision costs are, on average, larger than enrollment costs, but it also shows that there is significantly more heterogeneity in enrollment costs than in decision costs. As a result, 60 percent of enrollees have larger decision than enrollment costs (see Figure D.1 in the Online Appendix).



(a) Distribution of $\hat{\eta}^M_{it}$ and $\hat{\eta}^E_{it}$ (b) Distribution of $\hat{\eta}^M_{it}/\hat{\alpha}_{it}$ and $\hat{\eta}^E_{it}/\hat{\alpha}_{it}$

FIGURE 4. : Distributions of estimated switching costs

Note: The figure on the left reports the estimated distribution of both switching costs. The figure on the right reports the distribution of switching costs in US dollars. The figure on the right omits less than 0.1 percent of the simulated switching costs that have values smaller than -10.

Table 6 reports various statistics for the distribution of switching costs and

TABLE 5—: Estimated parameters of the structural model (switching costs)

		(1)	(2)	(3)
Decision cost	Constant	3.503	3.477	3.432
		(0.068)	(0.070)	(0.070)
	Age	0.015	0.015	0.016
		(0.002)	(0.002)	(0.002)
	Male	-0.02	-0.025	-0.021
		(0.034)	(0.034)	(0.034)
	Income	-0.004	-0.004	-0.004
		(0.001)	(4E-4)	(5E-4)
	Voluntary savings	-0.393	-0.396	-0.392
		(0.036)	(0.036)	(0.036)
	Regulation	0.520	0.539	0.531
		(0.037)	(0.038)	(0.038)
	σ_{η_D}		0.043	0.043
			(0.011)	(0.011)
Enrollment cost	Constant	1.175	2.551	2.538
		(0.084)	(0.279)	(0.253)
	Age	0.013	0.018	0.018
		(0.003)	(0.003)	(0.003)
	Male	0.003	0.026	0.023
		(0.041)	(0.045)	(0.045)
	Income	-0.005	-0.010	-0.010
		(0.001)	(0.001)	(0.001)
	Voluntary savings	0.164	0.113	0.112
		(0.043)	(0.048)	(0.048)
	Regulation	0.423	0.709	0.710
		(0.047)	(0.075)	(0.071)
	σ_{η_E}		1.988	1.962
			(0.211)	(0.194)
McFadden's Pseudo R^2		0.890	0.890	0.891
$-\left(\frac{1}{N}\mathcal{L}\right)$		0.267	0.266	0.266
Number of observations			350,660	

Note: Standard errors in parentheses. Specification (1) is the baseline. Specification (2) includes random coefficients. Specification (3) includes the interaction of PFA fixed effects and demographics. PFA fixed effects included in all specifications. The dependent variable is an indicator that is equal to one for the chosen PFA and zero for the rest. Estimation is via maximum likelihood in column 1 and simulated maximum likelihood in columns 2 and 3. Specifications (2) and (3) use 50 Halton draws per individual.

salaries in 2001 U.S. dollars. To compute switching costs in dollars, I divide estimated switching costs by the estimated marginal utility of income, $\hat{\alpha}_{it}$. The table shows that, on average, decision costs are \$37.5, while enrollment costs are \$30, both of which are higher than the average fee paid. This is a consequence of the low switching rate and relatively inelastic demand (low $\hat{\alpha}$), and is not uncommon (see, for example, Goettler and Clay, 2011, and Handel, 2013, among others). The table also shows that estimated switching costs are of the order of 10 percent of monthly salary.

TABLE 6—: Switching costs in monetary units

	Decision cost	Enrollment cost	Income
Mean	\$37.5	\$30.2	\$332.6
Median	\$34.2	\$27.1	\$220.0
Standard deviation	\$16.2	\$22.9	\$357.4

Note: Switching costs measured in dollars correspond to the ratio between estimated switching costs and $\hat{\alpha}_{it}$. Computation excludes the upper and lower 1 percent of the distributions to avoid issues with extreme values affecting computation of the mean and standard deviation. Income corresponds to monthly salary. All numbers are in 2001 dollars.

Estimated taste coefficients $\hat{\alpha}$ and $\hat{\beta}$ are reported in Table 7 and the distributions are presented in Figure 5. The table shows that older enrollees and males have a lower marginal utility of income than younger enrollees and females. At the same time, older enrollees derive a higher marginal utility from past performance than younger enrollees. This is consistent with older enrollees—who are closer to retirement—being more interested in maintaining higher balances than in minimizing fees, as new contributions represent a smaller fraction of their accumulated balances. Hence, older enrollees may try to keep their balances unaffected by returns in the short run, even if that is associated with paying a higher fee.

In terms of model fit, Table 8 reports actual and predicted average choice probabilities (market shares) for those PFAs with average market shares greater than 1 percent using three samples. The first three columns report choice probabilities according to the estimated parameters using the entire sample. The table shows that fit is quite good for the largest PFAs, but less good for smaller PFAs (though these represent less than 7 percent of the market). The second set of three columns uses the same set of estimates, but reports choice probabilities for initial choices only. These columns show that although overall fit is still good, it is not as good as that obtained when the whole sample is considered. This result is a consequence of some PFAs' having zero market share among initial choices, while the model predicts strictly positive probabilities for all of them. Finally, the last set of three columns reports choice probabilities among switchers.

Table 7—: Estimated parameters of the structural model (taste coefficients)

		(1)	(2)	(3)
Taste coefficient on income	Constant	4.1362	4.7118	3.1856
		(0.3295)	(0.3725)	(0.4571)
	Age	-0.0979	-0.1012	-0.0237
		(0.0097)	(0.0102)	(0.0135)
	Male	-0.7342	-0.8290	-0.9009
		(0.1602)	(0.1728)	(0.2012)
	σ_α		0.1611	0.1761
			(0.0675)	(0.0672)
Taste coefficient on returns	Constant	-0.045	-0.0478	-0.0318
		(0.0059)	(0.0063)	(0.0066)
	Age	0.0013	0.0014	0.0010
		(0.0002)	(0.0002)	(0.0002)
	Male	-0.0042	-0.0048	-0.0047
		(0.0030)	(0.00323)	(0.0033)
	Income	0.0005	0.0006	0.0005
		(4.6E-5)	(4.8E-5)	(4.7E-5)
	σ_β		0.0037	0.0044
			(0.0068)	(0.0070)
McFadden's Pseudo R^2		0.890	0.890	0.891
$-\left(\frac{1}{N}\mathcal{L}\right)$		0.267	0.266	0.266
Number of observations			350,660	

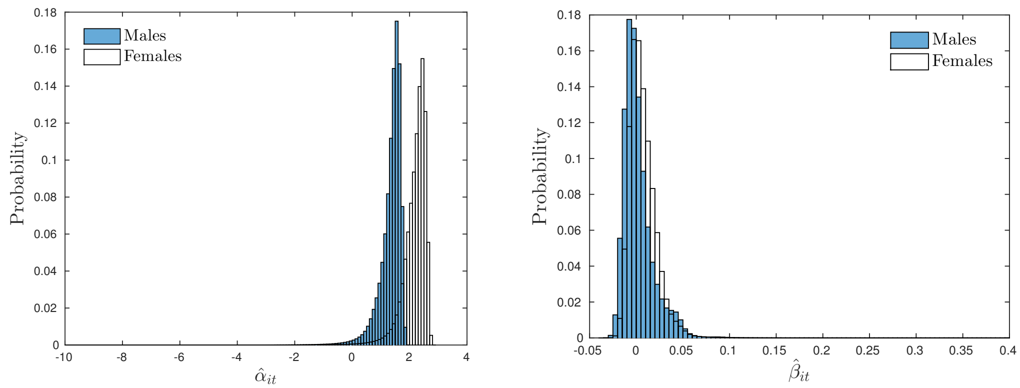
Note: Standard errors in parentheses. Specification (1) is the baseline. Specification (2) includes random coefficients. Specification (3) includes the interaction of PFA fixed effects and demographics. PFA fixed effects included in all specifications. The dependent variable is an indicator that is equal to one for the chosen PFA and zero for the rest. Estimation is via maximum likelihood in column 1 and simulated maximum likelihood in columns 2 and 3. Specifications (2) and (3) use 50 Halton draws per individual.

C. Robustness Analysis and Caveats

Here I present alternative specifications of the demand model to test its robustness. The results of these specifications are reported in Table D.1 and Table D.2 in the Online Appendix.

First, to consider how employment and income histories affect preferences, I include accumulated balances in the specification of switching costs and taste coefficients. Furthermore, as balances do not affect the fees an enrollee pays, including balances in the specification of α does not trigger the problems that including income would. The results, reported in the first column of Table D.1 and Table D.2, are similar to those reported in Table 5 and Table 7.

Second, in column 2 of Table D.1 and Table D.2, I include the time spent unenrolled before returning to the market as a determinant of decision costs. I include



(a) Distribution of $\hat{\alpha}_{it}$ (b) Distribution of $\hat{\beta}_{it}$

FIGURE 5. : Distributions of estimated taste coefficients

Note: The figures report the distribution of the marginal utility of income ($\hat{\alpha}$) and the distribution of preferences over the ranking of returns (β), by gender.

TABLE 8—: Model fit (choice probabilities, percent)

Share over All Choices			Share over Initial Choices			Share among Switchers		
PFA	Data	Predicted	PFA	Data	Predicted	PFA	Data	Predicted
20	36.6	33.3	20	37.1	27.3	20	24.6	23.9
12	24.5	24.0	22	22.5	16.7	12	21.4	20.9
22	18.1	20.3	12	20.5	18.9	22	10.9	15.7
23	7.1	6.9	23	4.7	8.0	7	9.4	7.7
7	5.9	4.6	19	3.4	7.0	23	8.9	7.8
19	2.1	4.8	6	3.1	3.8	19	5.7	6.5
13	1.5	1.2	24	1.9	7.5	6	3.6	8.8
24	1.3	3.0	7	1.7	6.2	1	3.3	5.2

Note: The table reports mean predicted choice probabilities for different samples. Predicted probabilities are computed using specification 6 of tables 5 and 7.

this variable because enrollees who have been away from the system for longer periods may have devoted more time to choosing a pension fund upon returning, in an attempt to compensate for those longer periods without contributions.[23] As before, the results are similar to those of the main specifications.

The third robustness check is related to the *i.i.d* assumption of ε. Though this assumption is common in the demand-estimation literature, it could happen that

[23]I do not include time unenrolled in the specification of enrollment costs as it would be zero for all existing enrollees and enrollment costs are sunk for returning enrollees.

in the setting studied in this paper, ε_{ijt} may be correlated over time. If this is the case, the model will attribute the induced inertia to switching costs rather than preferences, thereby overestimating switching costs. To account for this possibility, this section presents a specification that drops the *i.i.d* assumption and introduces autocorrelation in ε. This change implies that estimation cannot follow the approach associated with Equation 3. The details of how estimation is performed in this case are reported in the Online Appendix. The results, reported in column 3 of Table D.1 and Table D.2 show that allowing for autocorrelation does not significantly affect the estimated distributions of switching costs and preferences. Furthermore, the autocorrelation coefficient is not significant, suggesting that the cause of inertia in this setting is the existence of switching costs, rather than possible persistence in ε.

Fourth, in column 4 of Table D.1 and Table D.2, I study whether the results are robust to changing the way returns are measured. Specifically, I specify the indirect utility function as a function of the interaction of actual returns and the enrollee's account balance, in place of the ranking of returns. That is, I specify the indirect utility function as

$$u_{ijt} = \alpha_{it}(y_{it} - 0.1y_{it} - f_{jt} - p_{jt}\min\{y_{it}, \bar{y}_t\}) + (R_{jt} \cdot B_{it})'\beta_{it} - \eta_{it}\mathbf{1}[d_{it-1} \neq j] + \xi_j + \varepsilon_{ijt}.$$

where all variables are the same as the ones defined above, with the exception of R_{jt}, which now measures real returns rather than a position in the ranking of returns, and B_{it} which corresponds to the balance of individual i. The estimated parameters, reported in column 4 of Table D.1 and Table D.2, remain unchanged.

Finally, none of the approaches discussed above take into account that fees might be correlated with the unobserved component ξ. To address this concern, I follow Train (2009) and employ a control function approach using PFA age, the number of PFAs in the market, and PFA returns as instruments, in addition to unobserved PFA characteristics' not varying over time. The results, presented in column 5 of Table D.1 and Table D.2, show that the estimates do not change relative to those of the preferred specification.

V. Switching Costs, Choices, and Dynamic Price Competition

Having shown that switching costs are significant in the Chilean pension system, I now study how switching costs affect enrollees' behavior and competition among pension funds. Accordingly, I simulate several counterfactual scenarios in which enrollees must choose funds while facing all, some, or none of the switching costs present in the Chilean system.

I divide my discussion of the policy implications of switching costs into two parts. First, I study how enrollees' choices, fees paid, and accumulated balances change as switching costs decrease, conditional on the observed fee schedule. That is, in the first set of counterfactuals, PFAs are not allowed to re-optimize. This allows me to isolate the impact of switching costs on consumer behavior. Second, I study how competition among PFAs changes when switching costs change. That is, the second set of counterfactuals studies how switching costs affect dynamic competition among PFAs. Examining how switching costs affect dynamic competition requires additional assumptions, and so the two sets of exercises are presented separately below.

Finally, it is important to note that because participation is mandatory for those employed in the formal sector, reducing or eliminating switching can only affect the transfer that takes place between enrollees and PFAs. However, and though not considered in this paper, if reducing or eliminating switching costs decreases equilibrium fees, participation in the pension system may increase, because lower fees may increase participation in the formal labor market. Quantifying the magnitude of this benefit of decreasing switching costs is beyond the scope of this paper.

A. *Consumer Behavior and Switching Costs*

The counterfactual exercises presented here use the estimated parameters of the preferred specification (third column of Table 5 and Table 7). I simulate four counterfactuals using a 30 percent random sample of enrollees.[24] In the first counterfactual, I simulate the sequence of enrollee choices when they face all switching costs. Next, I simulate choices for the same sample when there are no enrollment costs. I then repeat the analysis, but eliminate decision costs. Finally, I simulate choices when there are no switching costs. Details of the procedures and an example are provided in the Online Appendix.

The results are presented in Table 9. The baseline when enrollees face all switching costs, shows that enrollees pay, on average, 5.89 percent more than if they were to choose the cheapest fund.[25] This suggests that individuals in the simulation tend to choose cheaper PFAs for their initial choice than what is recorded in the data, as they rarely switch PFAs when fees change (this is true in both the simulation and the data).

Next, when there are no enrollment costs, overpayment relative to the cheapest option decreases to 5.60 percent—a reduction of 0.29 percentage points. Accumulated balances remain unchanged.

The third counterfactual, which eliminates decision costs, decreases overpayment relative to the baseline but increases it relative to the case in which enrollment costs are eliminated. The overpayment rate is 5.77 percent, 0.12 percentage points lower than that of the base case.

Finally, the last row of Table 9 shows that overpayment increases when all switching costs are eliminated. Indeed, in the absence of switching costs, enrollees switch PFAs significantly more than in the previous cases (see Figure E.1 in the

[24] Results do not change significantly when using different random samples or increasing the sample size (though increasing the sample size requires reducing the number of simulated draws due to memory constraints).

[25] Because there is no outside option, I describe all fees relative to those charged by the cheapest fund.

Online Appendix for an example), but seem to choose funds based on non-fee attributes.

TABLE 9—: Overpayment and savings rates under counterfactual policies

	Overpayment rate		Mean final balance
Policy	Mean	Median	relative to base case
Base simulation	5.89	4.66	-
No enrollment cost	5.60	4.33	0.02
No decision cost	5.77	4.52	-0.02
No switching costs	6.01	4.81	0.03

Note: The table reports the mean and median overpayment rate and the mean balance as percentage of the base case for each of the policies under study. None of the mean differences reported in the last column is statistically different from zero. Also, differences in median savings relative to the base case are always zero.

In summary, eliminating some but not all switching costs results in enrollees' paying less than when they face none. This happens because PFAs are essentially differentiated products, and therefore when switching costs are eliminated, enrollees switch PFAs to pursue fund attributes that may or not be fees. Because enrollees care about these attributes, switching to relatively more expensive funds is justified. In addition, the higher switching rate may also be explained by the assumption that enrollees pursue myopic strategies. That is, if enrollees were to pursue dynamic strategies in an attempt to maximize utility at retirement, switching rates could be lower than those presented here.

B. Switching Costs and Dynamic Competition

I now turn to examination of how price competition changes when policy either reduces or eliminates switching costs. In this context, it is necessary to introduce a dynamic-competition model in which PFAs choose their fees to compete for enrollees who face different levels of switching costs. An important assumption of the model is that instead of requiring PFAs to keep track of each enrollee's status, I assume that PFAs make their decisions based on their shares of enrollees of each category. To compute these aggregate shares, it is necessary to integrate over the distribution of consumer preferences. That is,

$$s_{kt} = \int s_{ikt}(\boldsymbol{p}_t, \boldsymbol{X}_t, d_{it-1}, \boldsymbol{\xi}; \alpha_{it}, \beta_{it}, \boldsymbol{\eta}_{it}) dF(\alpha, \beta, \eta),$$

where $k \in \{\text{new}, \text{returning}, \text{existing}\}$, $\boldsymbol{p}_t$ is the vector of fees charged in period t, $\boldsymbol{X}$ is the vector of observable fund characteristics, d_{it-1} represents individual i's choice in the previous period, $\boldsymbol{\xi}$ is a vector containing all ξ_j, and $\boldsymbol{\eta}$ the inertia components that determine current market shares.

In this setting, if there are M_{kt} enrollees in status k at the beginning of period t, static profits of firm j are given by

$$\Pi_{jt}(\boldsymbol{s}_t, \boldsymbol{p}_t, \boldsymbol{X}_t) = \sum_k (p_{jt} - c_{jt}) M_{kt} s_{kt}(\boldsymbol{s}_{t-1}, \boldsymbol{p}_t, \boldsymbol{X}_t),$$

where s_{t-1} corresponds to the share vector at the end of the previous period, which contains information about each firm's share among consumers of each type. In the context of this paper, marginal costs are likely to be small, as the cost of the sales force or of the investment department does not depend directly on the marginal consumer. For this reason, I assume c_{jt} to be equal to zero for all PFAs.

Because switching costs introduce dynamics in the firm's problem, this can be represented as

$$V_j(\boldsymbol{s}_{t-1}, \boldsymbol{X}_t) = \max_{p_{jt}} \ \Pi_{jt}(\boldsymbol{s}_t, \boldsymbol{p}_t, \boldsymbol{X}_t) + \beta \boldsymbol{E}_t[V_j(\boldsymbol{s}_t, \boldsymbol{X}_{t+1})]. \tag{4}$$

The first-order necessary condition associated with Equation 4 is then

$$\frac{\partial \Pi_{jt}}{\partial p_{jt}} + \beta \left[\frac{\partial \boldsymbol{s_t}}{\partial p_{jt}}\right]' \boldsymbol{E}_t \left[\frac{\partial V_j(\boldsymbol{s}_t, \boldsymbol{X}_t)}{\partial \boldsymbol{s}_t}\right] = 0, \tag{5}$$

where $\left[\frac{\partial \boldsymbol{s_t}}{\partial p_{jt}}\right]'$ and $\boldsymbol{E}_t \left[\frac{\partial V_j(\boldsymbol{s}_t, \boldsymbol{X}_t)}{\partial \boldsymbol{s}_t}\right]$ are the two elements that follow from application of the chain rule.

In our setting, Equation 5 is one of many first-order conditions that have to be satisfied simultaneously (two per firm). In addition, the dimensions of the state space render the problem intractable. For these reasons, I introduce four assumptions so that I can numerically solve the dynamic problem. The first assumption reduces the firms' problem by assuming that instead of firms' having to choose both a fixed fee and a percentage fee, they choose only the percentage fee. This simplification is somehow natural, as the bulk of PFA revenues were generated by the percentage fee.[26]

The rest of the assumptions reduce the dimensions of the state space. Specifically, I assume that i) excess returns are always zero, meaning that all funds generate the same returns to their enrollees; ii) the share vector only consists of shares among returning and existing enrollees; and iii) I follow an approach inspired by Ifrach and Weintraub (2017) and Benkard, Jeziorski and Weintraub

[26]The fixed fee represents, on average, 8.8 percent of total revenues, with the median being 6.5 percent. Furthermore, PFAs' financial records show that the relevance of the fixed fee decreased over time. For example, in January 2000, for seven of the eight PFAs in the market, the percentage fee corresponded to between 93 and 100 percent of revenues generated by mandatory contributions. The exception was PFA Planvital, for which case the percentage fee generated 81 percent of revenues. Finally, the fixed fee was eliminated in 2008.

(2015), and assume that three of the PFAs in the data behave as strategic players and keep track of each other, while the rest are aggregated into a secondary option that is treated as nonstrategic. Though each of these assumptions is strong, they can be justified as follows. First, regarding excess returns, this is a natural simplification because returns vary little across PFAs. This assumption means that in period t, enrollees do not expect one PFA to generate higher returns than another in period $t + 1$. Second, eliminating new enrollees is not consequential, because though they certainly encourage competition, they are a small fraction of total enrollees in each period (and a decreasing fraction during the first decade of the system). For this reason, I assume that firms make their decisions based on the number of returning and existing enrollees.

Finally, the third assumption is probably the strongest, in that it reduces the number of firms that are strategic. However, in my data, three of the firms jointly held 90 percent of the market, which suggests that the assumption is less strong than it might appear. Overall, these assumptions allow me to numerically solve for equilibrium fees for each element of the state space. Subsection E.3 in the Online Appendix describes the algorithm used to compute these prices as well as the resources employed. I compute equilibrium prices for four cases that replicate the ones in the previous subsection. Results are presented in Table 10.

The first row of Table 10 reports that the mean expected fee for the base case (computed using the implied market shares given the equilibrium prices each firm charges) is 6.2 percent. This should serve as our baseline to evaluate the impact of the different policies I now introduce.[27]

Table 10, row 2, shows that when enrollment costs are eliminated, equilibrium fees decrease by 2.53 percentage points—a 41 percent reduction from the base case. On the other hand, when decision costs are eliminated (row 3), equilibrium fees decrease by 2.4 percentage-points—a 38 percent reduction. This suggests that, conditional on having to choose a single switching cost to eliminate, enrollees are better off when enrollment costs are eliminated. However, the results also show that eliminating all switching costs reduces equilibrium fees to 2.61 percent—the lowest among all options.[28]

In the two sets of counterfactuals presented above consumers are better off when enrollment costs rather than decision costs, are eliminated. However, the elimination of all switching costs reduces equilibrium fees the most—a result that was impossible to capture when firms were not allowed to re-optimize. This shows that because enrollees must participate in the market, switching costs result in significant transfers between enrollees and PFAs. Hence, policies that reduce or eliminate switching costs may not only reduce equilibrium fees significantly;

[27] The first scenario is meant to replicate what is observed in the data under the additional assumptions just described. For this reason, one should not expect this scenario to match the data, but simply to serve as a benchmark to evaluate the effectiveness of policies that either reduce or eliminate switching costs in this simpler environment.

[28] In is important to note that in this last case the game is static as the only source of dynamics is the existence of switching costs.

TABLE 10—: Equilibrium fees under counterfactual policies

Case	Mean and 95 percent CI
Base simulation	6.195
	[6.181,6.210]
No enrollment cost	3.666
	[3.660,3.671]
No decision cost	3.837
	[3.833,3.842]
No switching costs	2.607

Note: The table reports the mean expected fees and 95 percent confidence intervals for the different scenarios under study. Means and confidence intervals are computed using equilibrium fees obtained from $10,000$ random initial states.

though not studied here, they may also induce higher participation in the system. In this sense, this paper shows that identifying the nature of coexisting switching costs may have a significant impact by informing the design of policies that motivate consumers to be more active and intensify competition among firms.

VI. Conclusions

Although the existence of switching costs has been extensively documented in economics, little is known about their cause. Further, in situations with multiple switching costs, researchers have been unable to distinguish among them or identify their separate impacts on behavior. This paper quantifies the impact of two sources of inertia among enrollees in a defined-contribution pension system: the cost associated with analyzing financial information and choosing a pension fund, and the hassle cost in the form of a time-consuming bureaucratic process that enrollees must undergo when switching pension funds.

I exploit variation derived from changes in employment status to show that enrollees who return to the system after periods during which they did not save for retirement are four times more likely to switch pension funds than enrollees who have contributed continuously, which is consistent with decisions of returning enrollees' being affected by less sources of switching costs than decisions of enrollees that have contributed continuously. I then show that switching behavior does not seem to be explained by differences in preferences across groups of consumers, but instead by the nature of the switching costs they face.

To quantify the impact of switching costs on consumer behavior, I estimate demand for pension funds. The results show that switching costs are mostly determined by the cost of analyzing financial information and choosing a pension fund, while the rest is explained by the bureaucratic hassle cost associated with switching.

I then study how enrollee and firm decisions are affected by switching costs. I first examine how enrollees' behavior changes when the different switching costs are eliminated and firms are not allowed to re-optimize. Then, I study how switching costs affect dynamic price competition among pension funds. The results show that when firms are not allowed to re-optimize, enrollees become more active as switching costs decrease and switch pension funds more often. When firms are allowed to re-optimize, eliminating all switching costs intensifies competition more than when only one of the existing switching costs is eliminated. However, eliminating either source of switching costs alone also has a significant impact on equilibrium fees. This suggests that policies that eliminate specific sources of switching costs may result in significant savings for enrollees.

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