DYNAMIC INCENTIVES IN RETIREMENT EARNINGS-REPLACEMENT BENEFITS

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Abstract—We analyze dynamic incentives in pension systems created by the use of a small set of final years of earnings to compute benefits. Using social security records and household surveys from Uruguay, we show that self-employed workers and some employees of small firms respond to these incentives by increasing reported earnings in the benefit calculation window. We find evidence that suggests that these responses are explained by changes in earnings reporting and not in total earnings or labor supply. Back-of-the-envelope calculations indicate that this behavior increases the cost of pensions by about 0.2% of the GDP.

I. Introduction

DEFINED-BENEFIT pension systems typically use some measure of individual average earnings before retirement to calculate pension benefits. Many use a limited set of final years of earnings in this calculation. Using a small number of years of reported earnings can benefit retiring workers by forgiving periods of low earnings, unemployment, or informality, and is less demanding in terms of information than using full work histories. However, it creates dynamic incentives to increase reported earnings during the final years of work. The potential costs of these behavioral responses are well known to economists (Gruber, 2016), and have been present in news coverage and policy discussions. Still, the empirical evidence on responses to these incentives is remarkably scarce, with the exception of Fitzpatrick (2017), who documents this behavior for public education pension schemes in the United States. To the best of our knowledge, there is no evidence on whether behavioral responses to such incentives occur in private-sector pension systems, whether there are real or reported earnings changes, or whether self-employed workers and employees respond differently.

To fill this gap in the literature, this paper studies the responses of self-employed and employed workers to the dynamic incentives of a 10-year retirement-benefit calculation window in a countrywide social security system. We use matched employer-employee monthly social security administrative records from Uruguay between 1996 and 2016, where private-sector worker retirement benefits are calculated using the reported earnings in the final 10 years of work. Uruguay is of particular interest for several reasons. First, the 10-year window is representative of many developing countries and occupation-specific pension systems in developed countries.[1] Second, Uruguay has a well-established contributory retirement system, where these dynamic incentives affect all private-sector workers, and where the availability of high-quality records allows us to evaluate responses and their heterogeneity at a full-country scale.

This paper has three main goals related to dynamic incentives in the Uruguayan pension system. The first goal is to model and test the behavioral responses to these incentives from self-employed and employed workers. Self-employed workers are likely more able to take advantage of such incentives because they face lower costs of manipulating reported earnings (Slemrod, 1992). Similar arguments can be made for earnings of employees in small firms, which are more prone to informal employment and tax evasion than large firms (Kleven, Kreiner, & Saez, 2016; Kumler, Verhoogen, & Frías, 2020; Slemrod, 2019). The second goal is to assess the margins at which these responses operate, and tie them to the presence of under-reporting of earnings. The third goal is to understand the cost to the social security system generated by these responses.

Our empirical analysis evaluates whether workers respond to the incentives embedded in the social security system by strategically increasing reported earnings in the benefit calculation window (BCW), conditional on contributing to social security. We conduct event studies of reported earnings from five years before until eight years into the BCW, flexibly controlling for the effects of time, industry, and firm size. We compare the behavior of self-employed workers, employees of firms with fewer than ten workers, and employees of larger firms. We focus on male workers, whose minimum retirement age is fixed throughout the period and who are more likely to achieve eligibility for retirement.

We use two alternative measures of the start of the benefit calculation window. We start by using age 50, 10 years before the minimum retirement age. This strategy leverages the 10-year window for benefit calculation, combined with a minimum retirement age of 60 that exogenously and sharply increases the probability of retirement. Using this measure

Received for publication May 13, 2020. Revision accepted for publication January 18, 2022. Editor: Brian A. Jacob.

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We thank the editor and three anonymous referees for very constructive suggestions. We also thank Alvaro Forteza, Ariel Cancio, Gautam Gowrisankaran, Michal Kobielarz, Marco Manacorda, Martin Leites, Ianina Rossi, Jessamyn Schaller, Braulio Zelko, and multiple seminar participants for very helpful comments. We are very grateful to *Banco de Previsión Social* for providing the data for this study. The authors declare that they have no relevant or material financial interests that relate to the research described in this paper.

A supplemental appendix is available online at https://doi.org/10.1162/rest_a_01193.

[1] Five other Latin American countries use the final or best 10 years of earnings in benefit calculation, and another nine use shorter windows (OECD, 2014). In Africa, occupational public-sector schemes frequently use a short period of final years (Stewart & Yermo, 2009). In developed countries, short benefit calculation windows were common until recently and are still widely used in the public sector. Five European countries were using windows of 5 to 15 years in their national retirement systems until the mid-2000s (OECD, 2005). Many state and local civil servant pension plans in the United States use windows of three to five years, and similar periods are used in many European public-sector pension schemes.

The Review of Economics and Statistics, May 2024, 106(3): 762–777

https://doi.org/10.1162/rest_a_01193

allows us to conduct event studies with both the administrative data and pooled cross sections of household survey data, and compare reported and total earnings responses. Whereas in the administrative data individuals may have incentives to misreport their earnings, these incentives are not likely to be present in the household survey information. To gain better precision on the timing of the start of the BCW, we then leverage an auxiliary administrative data set on retirees to estimate a model of retirement age. Based on this model, we predict retirement in our main social security sample, and compute the start of each individual's BCW as ten years before predicted retirement age. Our main estimates of reported earnings responses in the social security data are based on event studies centered around this predicted start of the BCW.

A. *Preview of Results*

We find that self-employed workers start increasing reported earnings sharply after the start of the BCW. Within five years of the start of the BCW, reported earnings are about 6% higher than average reported earnings the year before its start. We also find suggestive evidence of increases in the reported earnings of employees of firms with fewer than ten workers, of only slightly smaller magnitude although less precisely estimated. Meanwhile, we find no evidence of changes in reported earnings for employees of larger firms. In household surveys, we find no evidence of changes in hours of work or total earnings for self-employed or employed workers. The comparison between event studies in administrative and household survey data, along with the analysis of other margins of response, suggests that responses to the dynamic incentives of the BCW are explained by changes in earnings under-reporting behavior.

Potential challenges to identification could come from changes in productivity or hours of work with age. Two sets of findings rule out this concern. First, we find that reported earnings of all workers are flat in the years leading to the beginning of the BCW. Thus, any confounding increases in productivity or labor supply would have to start just after that, and only for self-employed and small-firm workers. Second, our analysis of earnings and hours of work in household surveys provides direct evidence that neither increases after the start of the BCW. We also provide several robustness checks and heterogeneity analyses, including sample attrition, estimates of models with individual fixed effects, dynamic difference-in-differences with large-firm workers as control group, a placebo check using employees of publicly owned firms, and estimates of responses for individuals with stable reporting or employment.

Finally, we provide back-of-the-envelope calculations of the implied cost of these responses to the social security system, only considering responses among the self-employed. We quantify costs under three scenarios of counterfactual reported earnings in the absence of these responses. First, we assume that the responses are entirely driven by over-reporting in the BCW. In the second scenario, we assume that they are entirely driven by under-reporting before the BCW. The third scenario assumes that counterfactual reported earnings would be the average of the earnings reported before and after the start of the BCW. The estimated increase in net expenditures on retirement benefits is between 3.3% and 4.2% per self-employed worker. This represents about 0.2% of Uruguay's GDP, which is roughly equivalent to annual government spending on early child care (MEC, 2019).

B. *Related Literature*

The contributions of this paper to the literature are twofold. First, to the best of our knowledge it is the first paper to document that reported earnings to social security respond to the dynamic incentives of short benefit calculation windows in a countrywide private-sector retirement system. This adds to a recent literature that has started to document responses to the dynamic incentives embedded in the computation of earnings-replacement benefits in pension systems (Fitzpatrick, 2017; Mannino & Cooperman, 2015). So far, this literature has focused on public service occupations with formal and very stable employment relationships and collective bargaining between employers and employees. In that specific context, employers and employees may be more able and willing to cooperate to defer compensation until after retirement if that can increase total rents due to political economy considerations (Glaeser & Ponzetto, 2014), even if workers' valuation of an increase in retirement benefits is well below its cost (Fitzpatrick, 2015). We show that private-sector workers also respond to the dynamic incentives of a retirement system with a limited benefit calculation window.

Second, this paper adds evidence to a large literature on tax evasion, by providing evidence that self-employed workers and small-firm employees respond to pension benefit incentives, and that responses most likely occur through changes in reporting behavior. Previous empirical work has found stronger responses to incentives created by tax and transfer schedules among self-employed and small-firm workers (Kleven et al., 2011; Kleven et al., 2016; Kuka, 2014; Kumler et al., 2020; Saez, 2010). More broadly, this paper contributes to a growing literature that studies how informality, reported earnings, and tax evasion respond to taxes and transfers in developing countries (Bergolo & Cruces, 2014; Best et al., 2015; Gerard & Gonzaga, 2021).[2]

[2]This paper also relates to the broad literature that analyzes the implications of retirement benefits in the design of pension systems, including the role of income replacement on the retirement decision (Asch, Haider, & Zissimopoulos, 2005; Biasi, 2019; Manoli & Weber, 2016; Van Soest & Vonkova, 2014); the interaction between retirement and labor supply (Blundell, French, & Tetlow, 2016, offer an excellent summary); and dynamic optimization problems in labor supply responses (Baker & Benjamin, 1999; Friedberg, 2000; Gelber et al., 2020; Haider & Loughran, 2008; Song & Manchester, 2007). We contribute to this literature by providing evidence of responses to dynamic incentives, studying an alternative

II. Institutional Background

In this section, we summarize the key elements of the institutional setting that are relevant for our analysis. A more detailed account of the setting is provided in online appendix A. We focus on self-employed workers (business owners of sole proprietorship firms) and private-sector employees in Uruguay who contribute to social security. Firm owners are required by law to register the firm with Uruguay's social security administration, *Banco de Previsión Social* (BPS), and register all employees and, when it corresponds, the owner's self-employment. This gives employees and self-employed workers obligations in terms of contributions, as well as the right to social insurance coverage.[3] To qualify for retirement, workers must be at least 60 years old and have contributed to social security for a minimum of 30 years of service. We focus on men, because women are less likely to achieve the required years of service by age 60, and their minimum retirement age was lower for some years of the sample period.

The social security system was reformed in 1996, when the current two-pillar pension regime was established. The *two-pillar* regime has a defined-benefit pillar, administered by BPS, and a defined-contribution individual savings accounts pillar. To limit negative impacts of the reform on workers closer to retirement, the reform also defined a *transition* regime for workers aged 40 or more by April 1996, which only has the defined-benefit pillar. The first 15 cohorts in our sample are subject to this transition regime, while the last 15 cohorts are subject to the new two-pillar regime. Even in the two-pillar regime, most workers can expect to receive a larger share of their retirement benefits from the defined-benefit pillar, especially those with lower lifetime reported earnings.

Retirement benefits from defined-benefit pensions (in both the transition regime and the defined-benefit pillar of the two-pillar regime) are calculated as the product of the reference average pre-retirement earnings and a replacement rate. Replacement rates are incremental with age at retirement and years of service, with a minimum of 45% (at age 60 with 30 years of service) and a maximum of 82.5% (at age 70 with 50 years of service). There is also a minimum and a maximum retirement benefit. Reference average pre-retirement earnings are calculated as the average of reported earnings of the final 10 years of work, as long as this does not exceed the average of the highest 20 years of earnings plus 5%. If the average earnings of the final 10 years of work is less than the average of the highest 20 years of earnings, the latter is used. This design of the benefit calculation introduces an incentive to report higher earnings in the 10 final years of work to increase the reference average pre-retirement earnings and thus retirement benefits.

Self-employed workers are subject to the same social security system as employees, although with some specific regulations. While employee earnings are reported to BPS by employers, who are responsible for paying employer and employee payroll taxes, self-employed workers must self-report their own earnings. Self-employed workers report earnings to BPS for the sole purposes of social security contributions and the calculation of benefits from social security and social insurance programs administered by BPS. Earnings are reported by choosing a category within a fixed scale, where each category is a multiple of a standard contribution base (adjusted yearly for inflation). The first of a total of ten categories in the scale is the minimum contribution base (MCB). Self-employed workers choose which category to report, within regulations that establish the minimum and maximum levels of earnings that can be reported, as well as how often the level of reported earnings can be changed, and by how much.

Self-employed individuals in two-pillar cohorts have more flexibility in the level of earnings they can report and how frequently and by how much they can change them. Workers in transition cohorts report earnings within standard categories, can only change reported earnings after spending 3 years in a category, and can only move up one category at a time. Workers in two-pillar cohorts can report any level of earnings above the MCB, including earnings above the last category. Self-employed workers who have employees have a higher minimum contribution level and cannot report earnings that are below the highest paid employee in the firm.

We take advantage of the features of the institutional setting to study responses of self-employed and employed workers to the benefit calculation window. The key elements of the setting for our empirical strategy are the 10-year benefit calculation window and the minimum retirement age of 60. The regulatory differences between groups of self-employed workers provide variation that allows us to study whether different types of restrictions to their earnings reporting are effective at mitigating responses.

III. Conceptual Framework

We develop a simple and stylized model of individual decisions in the presence of pension benefits that are a function of reported earnings in the final years of work. The model has three time periods; in periods 1 and 2 individuals work, and in period 3 they are retired. In period 1, reported earnings do not affect retirement benefits; period 2 earnings are in the retirement-benefit calculation window. We assume that the individual retires at the end of period 2, and this decision is exogenous to the model.[4]

source of potential effects on labor supply, and highlighting a new mechanism that could potentially affect retirement timing choices.

[3] Some groups of workers are exempted from the obligation to register with BPS. These are workers in occupation-specific retirement schemes that are administered separately: bank employees, notaries, and university graduates in independent exercise of liberal professions. In this paper we focus on workers covered by the general private-sector social security scheme, excluding public-sector workers. We also exclude rural and construction workers, who are subject to specific contribution regulations.

[4] Retirement age may also be endogenous to the design of the BCW. Modeling retirement age responses is beyond the scope of this paper. Our

We consider an individual i deciding the level of earnings to report to the social security administration in each of the two working periods (w_{i1} and w_{i2}). In each period of work, the individual is endowed with exogenous pretax total earnings Z_i. The model could be extended to account for endogenous labor supply; however, we discuss evidence in section VC that suggests that labor supply is not a relevant margin of response in our setting. A person reporting w_{it} earnings pays τw_{it} in payroll taxes, with τ being the tax rate over reported earnings. There is a minimum contribution base of $\bar{w}$ that applies in both periods ($w_{i1}, w_{i2} \geq \bar{w}$ for all i). If reported earnings are below true earnings, there is an expected cost $f(Z_i, w_{it})$, which we assume to be weakly increasing in total earnings ($\partial f(\cdot)/\partial Z_i \geq 0$), and weakly decreasing in reported earnings ($\partial f(\cdot)/\partial w_{it} \leq 0$). We allow the functional form to be general in how it depends on these two arguments.

Under these assumptions, expected consumption in periods 1 and 2 is given by $c_{it} = Z_i - \tau w_{it} - f(Z_i, w_{it})$. Consumption in the third period is given by $c_{i3} = p(w_{i2})$, where $p(w_{i2})$ is income from retirement benefits. Pension benefits are assumed to be weakly increasing in reported earnings in period 2 ($\partial p(\cdot)/\partial w_{i2} \geq 0$).[5] We assume individuals use the interest rate (r) to discount the future, and there is no saving or borrowing except through the pension system.

Utility in each period is given by $U(c_{it})$, with $\partial U(\cdot)/\partial c_{it} > 0$, $\partial^2 U(\cdot)/\partial c_{it}^2 < 0$. The individual discounted utility maximization problem is

$$\max_{w_{i1}, w_{i2}} V(w_{i1}, w_{i2}, Z_i) = \sum_{t=1}^{3} \left(\frac{1}{1+r}\right)^{t-1} U(c_{it})$$

subject to $\tau w_{it} + f(Z_i, w_{it}) \leq Z_i$ and $w_{it} \geq \bar{w}$ for all i and $t = 1, 2$. Given that there is no borrowing or saving, we assume that the sum of the taxes and the penalization is always strictly less than the total earnings that the individual receives each period ($\tau w_{it} + f(Z_i, w_{it}) < Z_i$). The first-order conditions of this problem provide the optimal solution for reported earnings in periods 1 and 2.

A. *First Period*

Optimal reported earnings in period 1 (w_{i1}^*) are given by

$$\tau = -\frac{\partial f(Z_i, w_{i1})}{\partial w_{i1}}$$

with $w_{i1}^* > \bar{w}$, or $w_{i1}^* = \bar{w}$. Intuitively, the optimal level of reported earnings is the one that equalizes the marginal cost of reporting one dollar (τ) in period 1 to its marginal benefit, expressed in terms of the expected reduction in the cost of misreporting earnings. If the optimal reported earnings are lower than or equal to $\bar{w}$, the individual reports $\bar{w}$. Thus, in period 1, optimal reported earnings are independent of expected pension benefits.

This optimal condition also highlights that heterogeneity in total earnings (Z_i) introduces heterogeneity in optimal reported earnings in period 1. Individuals with low earnings and low probability of enforcement will report the minimum contribution base. The sign of the cross partial derivative of the evasion cost function with respect to total and reported earnings ($\partial^2 f(\cdot)/\partial w \partial Z$) determines where in the income distribution individuals gain more from reporting an additional dollar of earnings. If this cross derivative is positive, higher-earnings workers face a steeper penalty as a function of reported earnings, and benefit more from reporting an additional dollar. Higher-income workers may face a steeper penalty function if, for example, the regulator monitors higher-income individuals with relatively higher probability when they report low earnings, but monitors them with a similar probability to lower-income workers when they report high earnings.

B. *Second Period*

Optimal reported earnings in the second period (w_{i2}^*) are given by

$$\tau = -\frac{\partial f(Z_i, w_{i2})}{\partial w_2} + \frac{1}{1+r}\frac{\partial p(w_{i2})}{\partial w_{i2}} \frac{\frac{\partial U(p(w_{i2}))}{\partial w_{i2}}}{\frac{\partial U(Z_i - \tau w_{i2} - f(Z_i, w_{i2}))}{\partial w_{i2}}}$$

with $w_{i2}^* > \bar{w}$, or $w_{i2}^* = \bar{w}$. Intuitively, during the BCW the optimal level of reported earnings is the one that equalizes the marginal cost of reporting an extra dollar to its marginal benefit, which now includes, in addition to the expected reduction in the cost of misreporting earnings, the discounted marginal utility benefit from increased pension benefits in period 3.

These conditions imply that reported earnings in period 2 are weakly higher than in period 1, and that there are two margins of response. In the extensive margin, individuals who report $\bar{w}$ in period 1, but are close to the margin, will report more than $\bar{w}$ in period 2. In the intensive margin, those reporting more than $\bar{w}$ in period 1 will report higher earnings in period 2 if they have positive marginal utility from the increase in pension benefits (since p is a weakly increasing function of w_{i2} and marginal utility is increasing in consumption). However, not everyone will increase reported earnings in the second period. As in the first period, some individuals will find it optimal to report $\bar{w}$. Others may derive no additional utility from increasing reported earnings in period 2, because they may be at a point where the pension benefit function is flat.[6]

identification strategy relies on the minimum retirement age and predicting retirement with predetermined observables.

[5] In the empirical setting, pension benefits may also depend on reported earnings in period 1 because of the computation method discussed in section II. However, these incentives are less salient to individuals and their effects more discounted over time. We focus on the incentives in the final 10 years of work. In this sense, we can interpret the model as the additional incentives to increase reported earnings in period 2.

[6] In the institutional setting of our empirical application, there is a minimum and a maximum retirement benefit. Individuals with very low optimal

Additionally, the model can be extended to reflect other potential sources of heterogeneity, such as costs of evasion, $f(Z, w)$, including for example perceptions about the probability of enforcement, and time preferences (represented by r). Another possible extension is to allow for a cost of changing reported earnings; this could serve to illustrate administrative or regulatory costs of adjustment for self-employed workers, which we introduce in our empirical application, or to represent in a simplified manner the costs of adjustment of reported earnings faced by dependent workers. Finally, the benefit from increased reported earnings in the BCW can depend on the individual's history of reported earnings. In particular, workers strongly attached to formal employment (those with high density of contribution) are more likely to satisfy the requirements for retirement.

For employed workers, a complete model would require addressing the dynamic bargaining between workers and employers. Employee earnings are reported by their employers. Firms have incentives to under-report employee earnings to lower their payroll tax bill, but run the risk of penalties if the under-reporting is revealed by an employee. They can mitigate this risk by making collusive agreements with their employees in which they share the surplus of tax evasion through informal (under-the-table) payments. Such collusion is less likely to be sustainable in large firms (Kleven et al., 2016). Consequently, the degree of under-reporting of employee earnings should be lower than that of self-employed workers and decreasing in firm size. This is consistent with previous empirical evidence (Kleven et al., 2011; Kumler et al., 2020), and with the descriptive evidence discussed in section IVB. Thus, the margin to adjust reported earnings is smaller for employees of large firms. In online appendix D.9, we discuss the potential collusion agreements between employers and employees and effects at the firm level. Overall, we only expect changes in reported earnings of employees in small firms, which are more prone to misreporting, and especially for workers with strong formal employment attachment and stable employment relationships.

IV. Data and Descriptive Statistics

A. Data Sources

Social security records (main analysis sample). Our main empirical analysis is based on matched employer-employee administrative records from Uruguay's social security system. The full data set contains the work histories between April 1996 and March 2016 of a random sample of about 300,000 workers (approximately 15% of the Uruguayan labor force) who were registered in social security at least 1 month during the period. The data were provided by Uruguay's social security administration (BPS). The structure of the data is an unbalanced panel of workers, with information on monthly reported earnings, contribution scheme, gender, date of birth, job start and end dates, type of remuneration, work schedule, firm ownership type, number of employees, and five-digit industry codes.

We keep observations of male workers contributing to the main social security contribution scheme, which includes workers in private-sector firms in the secondary and tertiary sectors. We exclude public-sector workers,[7] and workers subject to specific rural and construction social security contribution schemes. We restrict the sample to workers whose type of remuneration is monthly salary, which is the most frequent and reliable form of reporting earnings. In light of our research design, we keep workers born between April 1941 and March 1971 (30 birth-year cohorts). In regressions that use age 50 as the start of the BCW, we keep observations between ages 45 and 57. In the analysis where we predict the start of the BCW at the individual level, we extend the data set to include all observations up to 8 years after the predicted start of the BCW. These restrictions imply that each cohort is potentially observed for at least 1 year, and does not enter the sample after age 55. We restrict the sample to workers who are either employed or self-employed and who report positive earnings for at least 6 months while in sample. Self-employed workers are identified as working owners of sole proprietorship firms. We adjust earnings to thousands of Uruguayan pesos (UY$) of December 2015. For reference, in December 2015 the exchange rate was about UY$30 per USD. We drop observations with reported earnings below the 5th or above the 95th percentiles, and observations with missing data on any of the variables we use in the analysis. See online appendix B.1 for more details.

Our main analysis sample has 5,091 self-employed workers (289,603 observations at ages 45 to 57) and 12,542 employed workers (654,517 observations). The average self-employed (employed) worker has a contribution density of 71% (67%).[8] We study employees of small (fewer than ten workers) and larger firms separately. To do this, we stratify the sample of employed workers on the basis of the size of the firm where they are employed when we first observe them in this sample. This leaves us with 4,355 individuals working in firms with fewer than ten workers when first observed (218,556 observations).

Household surveys. We supplement our analysis with information from annual cross-sectional household surveys (*Encuesta Continua de Hogares*, ECH), conducted by the National Statistics Institute (INE). From these surveys,

reported earnings in period 1 may still be at the minimum pension benefit even if they increase reported earnings in period 2, while individuals with high earnings may hit the benefit ceiling with the optimal reported earnings of period 1.

[7] We use the data on workers of publicly owned firms in a placebo test.

[8] Note that, given the start and end dates of the data set, and our cohort and age restrictions, not all individuals are potentially observed for the whole period between ages 45 and 57. We define the contribution density as the number of months of contribution to social security divided by the maximum number of months an individual could potentially be in our sample given his birth cohort.

TABLE 1.—DESCRIPTIVE STATISTICS

	Social security data				Household surveys			
	Self-employed		Employed		Self-employed		Employed	
Reported earnings (UY$1,000)	9.50	(5.58)	26.51	(18.42)	31.03	(17.04)	27.50	(15.26)
Subsample small firm[a]			17.31	(12.32)			21.16	(11.22)
Reported < full salary[b]							0.06	(0.24)
Subsample small firm[c]							0.11	(0.32)
Reported earnings/MCB[d]	1.29	(0.71)						
Reported earnings = MCB	0.55	(0.50)						
Reporting density[e]	0.71	(0.28)	0.67	(0.27)				
Hours of work/week	40.82	(2.64)	41.31	(6.41)	49.27	(13.55)	47.74	(9.36)
Age	51.13	(3.70)	50.92	(3.68)	50.82	(3.66)	50.65	(3.66)
Firm with 0 employees	0.65	(0.48)	0.00	(0.00)	0.47	(0.50)	0.00	(0.00)
Firm with 1–5 employees	0.29	(0.45)	0.19	(0.39)	0.43	(0.50)	0.13	(0.34)
Firm with 5–9 employees	0.05	(0.21)	0.13	(0.34)	0.11	(0.31)	0.12	(0.33)
Firm with 10+ employees	0.02	(0.14)	0.68	(0.47)	0.00	(0.02)	0.75	(0.44)
Manufacturing	0.19	(0.39)	0.16	(0.37)	0.17	(0.38)	0.30	(0.46)
Retail, hospitality	0.37	(0.48)	0.32	(0.47)	0.45	(0.50)	0.23	(0.42)
Transportation, communication, energy	0.19	(0.39)	0.12	(0.32)	0.15	(0.36)	0.21	(0.41)
Services, other	0.26	(0.44)	0.39	(0.49)	0.22	(0.42)	0.25	(0.43)
Observations	289,603		654,517		9,384		28,855	
Individuals	5,091		12,542		9,384		28,855	

Sample mean (s.d.) from the main administrative sample and household surveys 1996–2016. Details in section IV. [a]Reported earnings in small firms (218,556 and 7,439 obs.). [b,c]Available since 2006 (22,628 and 5,731 obs.). [d]Reported earnings divided by minimum contribution base. [e]Months with reported earnings divided by maximum potential months in the sample.

which are representative of the working age population, we obtain earnings and usual working hours. We identify formal workers as those covered by social security. Self-employed workers are those who report working for themselves or working as business owners as their main occupation. We restrict the samples closely following the procedures used in the social security data. We drop workers in the primary sector, construction, and the public sector, and those who contribute to other occupational pension schemes (independent liberal professionals and notaries). We keep male workers born between 1941 and 1971, and observations between ages 45 and 57. We adjust earnings to thousands of Uruguayan pesos (UY$) of December 2015. We drop observations with reported earnings below the 5th or above the 95th percentile, and observations with missing data on any of the variables we use in the empirical analysis. The final sample has 9,384 observations of self-employed workers and 28,855 observations of employees. Among the latter, 7,439 observations are in firms with fewer than ten workers. More details are provided in online appendix B.2.

Measures of earnings. The social security records report gross earnings before worker social security contributions. In household surveys, individuals report net monthly earnings, after taxes and contributions. To obtain a comparable measure of monthly gross earnings in the household surveys, we impute social security contributions and add them to net earnings. More details are provided in online appendix B.3.

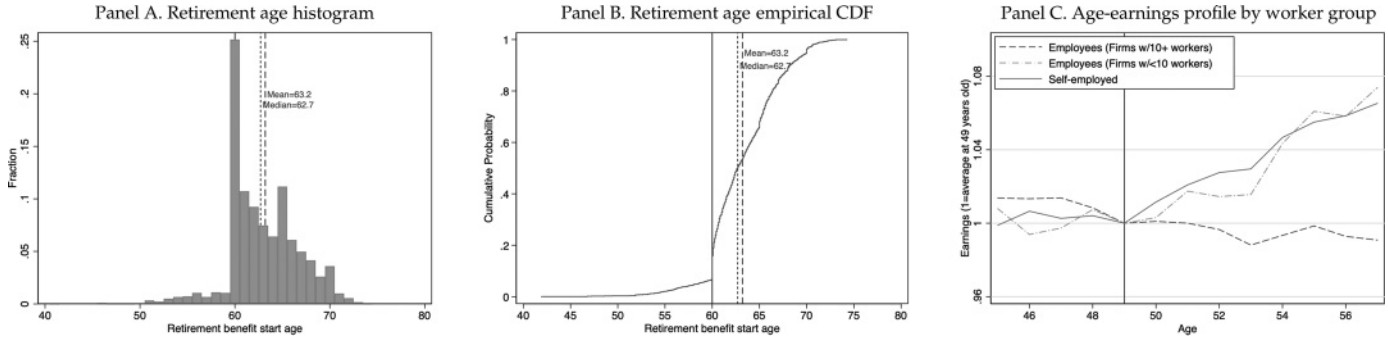
Supplementary social security data sets. We use two auxiliary administrative data sets from BPS to conduct supplemental analyses. The *retirement sample* has retirement and work history records for a sample of retirees receiving benefits between December 2001 and February 2016. Details about this sample are provided in appendix E.1. The *firms sample* has administrative records of a random sample of firms operating between 1996 and 2013. The data set follows each firm during that period, including worker-level records while workers are employed at the firm. More details of the data set and the empirical analysis are provided in online appendix D.9.

B. Descriptive Statistics

Table 1 reports summary statistics for the main analysis samples from the social security records and household surveys at ages 45 to 57. The first two columns report summary statistics for self-employed workers in the social security records. Self-employed individuals mostly have very small firms, with 65% having no employees. Their average reported earnings are UY$9,500 per month, equivalent to about US$316, or 1.3 times the MCB. In 55% of observations, they report earnings exactly equal to the MCB. The third and fourth columns show summary statistics of social security records for employees. The mean of reported earnings in this sample is UY$26,510 (US$883), 2.8 times the average reported earnings of self-employed workers. Employees are more likely to work in larger firms and in services than self-employed workers. Those working in small firms (with fewer than ten employees) have lower average reported earnings.

In the last four columns of table 1, we present summary statistics for self-employed and employed workers in the household surveys. For the full sample of employees, average reported earnings are very similar to those reported in the administrative records. For employees working in small

FIGURE 1.—DESCRIPTIVES OF RETIREMENT AGE AND AVERAGE EARNINGS BY AGE



(A) Histogram of discrete age in years at the time of retirement. (B) Empirical cumulative distribution function of continuous age at retirement. Data from the *retirement sample* of social security records for workers who are observed collecting retirement benefits between 2001 and 2015 and are age 65 or above by the end of the panel. (C) Average monthly earnings reported to social security of self-employed and employed workers by age, normalized to 1 at age 49. The sample includes all observations for self-employed and employed workers ages 45 to 57, reporting earnings for at least 1 month between January 2005 and March 2016.

firms, however, average earnings reported in household surveys are 22% higher than those reported to social security. Since 2006, household surveys directly ask employees if their earnings are under-reported to social security; 11% of employees of small firms respond that their earnings are under-reported. The contrast in reported earnings between the two data sources is even more striking for the self-employed. In household surveys, self-employed workers report earnings over three times the average earnings reported to social security. This suggests that earnings of self-employed workers are substantially under-reported to the social security administration, as are, to a lesser extent, those of employees of small firms. Self-employed workers also report having more employees in household surveys than what we observe in social security records, which could be related to informal employment.

Figure 1 shows the histogram (panel A) and cumulative distribution function (panel B) of retirement age from the *retirement sample*. We identify retirement as the first month the worker receives contributory retirement benefits.[9] There is a very sharp jump in the probability of retirement at the minimum retirement age: 25% of the sample starts receiving benefits at age 60. Less than 10% retire before this age, which is almost completely explained by early retirement due to disability. In the years that follow the minimum retirement age, the cumulative percentage of workers that have retired increases rapidly, with over 50% of workers retiring before the 63rd birthday. The average retirement age is 63.2. These observations, combined with the way retirement benefits are calculated, motivate our use of age 50 as the threshold age at which we expect to start observing increases in earnings reported to social security in order to increase future retirement benefits.

[9]We also observe the date when each worker initiates the retirement paperwork, which is on average 4 months before they start receiving benefits. In our main sample, we observe a proxy for retirement for a small group of workers—based on firms' reported reason for a worker's discharge from payroll—which has a very similar distribution.

C. *Descriptive Evidence*

Before presenting our regression framework, we begin by motivating our analysis with some descriptive evidence. If the predictions of our model are correct, we expect average reported self-employment earnings to start increasing at age 50, the minimum age for starting the 10-year BCW. In figure 1, panel C, we plot average reported earnings by age of self-employed and employed workers, in the years around their 50th birthday. We distinguish between employees of very small firms (fewer than ten workers), which are more prone to informality and under-reporting, and larger firms.[10] The reported earnings of all workers are stable in the years leading to age 50, and those of large-firm employees continue on the same (slightly decreasing) path after age 50. This pattern is consistent with the patterns found by Lagakos et al. (2018) for potential experience-earnings profiles in several high and upper-middle income countries, including Uruguay. In contrast, reported earnings of self-employed workers and small-firm employees start increasing rapidly at age 50. This evidence is consistent with self-employed and small-firm workers responding to the incentives of the retirement system.

In addition, the comparison of average reported earnings for each worker group across data sources suggests that these increases in reported earnings represent a decrease in under-reporting. Online appendix table B3 shows that, for large-firm employees, mean earnings are very similar across the two data sources for all age groups. In contrast, mean earnings are substantially lower in the administrative data than in the household survey data for small-firm employees and self-employed workers. The degree of under-reporting in the administrative data decreases after age 50 for these two groups, in a magnitude that is very close to the estimated reported earnings responses from the regression analysis presented in the next section, especially for the self-employed.

[10]In this purely descriptive analysis, we do not control for any potential confounders such as time effects. Thus, to avoid the recession that started in 1999 and culminated with a deep financial crisis in 2002 and a sharp recovery in the 2 years that followed, we exclude pre-2004 data.

V. Identifying the Responses of Reported Earnings to the Benefit Calculation Window

A. Empirical Approach

We implement dynamic event studies of reported earnings around the start of the benefit calculation window, controlling for potential confounding factors, including time, sector, and firm-size effects. We estimate the following model:

$$Y_{it} = \sum_{b=-5}^{-2} \mu_b 1[\text{Years Window}_{it} = b] + \sum_{a=0}^{7} \mu_a 1[\text{Years Window} = a] + \delta_t + \gamma X_{it} + \varepsilon_{it}. \quad (1)$$

Y_{it} is the outcome variable for individual i in period t; our main outcome variable is reported earnings to social security, expressed as the ratio over time -1 average reported earnings, so that estimates can be interpreted as percent changes. Years Window$_{it}$ is the number of years relative to the start of the 10-year benefit calculation window, which we define below. Parameters μ_b and μ_a represent changes in reported earnings before and after the start of the BCW, respectively, relative to the omitted event time (-1). We control flexibly for time effects by including dummy variables for each month-year (δ_t). We also control for firm size categories and economic sector (two-digit ISIC codes), and their interactions with calendar year. We report robust standard errors clustered by individual.

To conduct the analysis, we stratify the sample in three groups: self-employed workers, employees of small firms (with fewer than ten employees), and employees of larger firms (ten or more employees). The latter two groups are defined according to firm size when the individual is first observed in the sample. We choose ten workers as the threshold because very small firms are more likely to have closer management-worker relationships, to be sole-ownership firms, and to be exposed to less stringent tax enforcement.[11] There is also a more practical reason for this threshold: in the household survey, respondents are asked to report if they work in a firm larger or smaller than ten employees, and the exact number of employees is only asked for firms with fewer than ten workers. Thus, breaking up firm size at ten workers allows us to conduct the same analysis in the household survey data as in the administrative records.

We use two alternative measures for the timing of the start of the BCW. We begin by using age 50, 10 years before minimum retirement age. We only observe actual retirement age for the older cohorts in our main analysis sample, and we do not observe it in the household survey data. Using age 50 as a benchmark allows us to compare social security reported earnings responses to the total earnings responses estimated with household survey data. While retirement age can be endogenous to changes in earnings during the final years of work, reaching the minimum retirement age exogenously increases retirement likelihood. As shown in figure 1, a quarter of retired workers start receiving retirement benefits at exactly the minimum retirement age. In addition, in a labor market with substantial employment and earnings volatility, actual retirement age can be hard to predict for the worker, and the minimum retirement age could serve as a focal point when planning contributions and retirement. For this analysis, we limit the sample to observations between ages 45 and 57, up to 5 years before and 8 years after the start of the BCW.

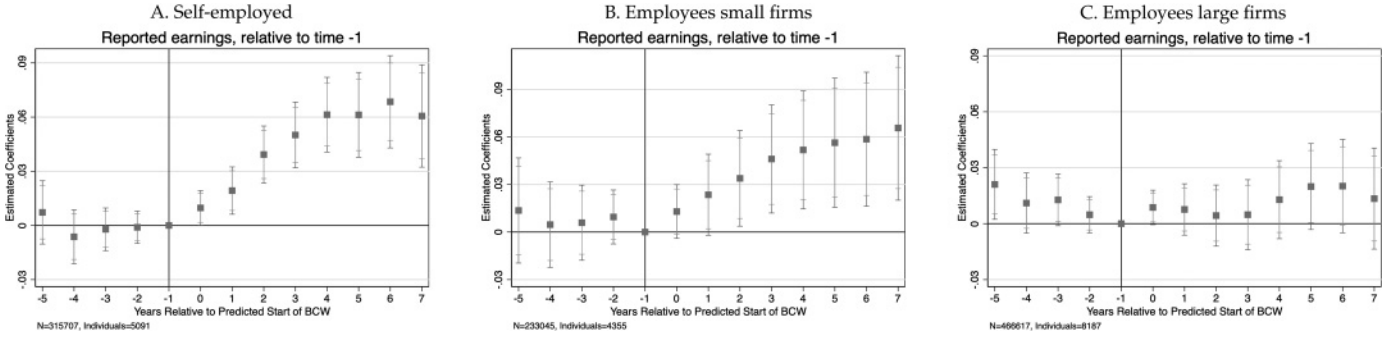
Because many workers retire later than the minimum retirement age, using age 50 provides a noisy measure of the start of the BCW. We construct a more precise measure of the start of the BCW based on the predicted retirement age for each worker in our main analysis sample. To this purpose, we use supplementary data from the retirement sample to estimate an ordered probit model of retirement age as a function of variables constructed with data observed up to age 49, including sector, type of employment, firm size, reporting above the MCB (for self-employed), and summary measures of reporting history. These variables are observed in both the retirement sample and the main analysis sample, so we can use the estimated model to predict the probabilities of retirement by age for individuals in our main analysis sample. The details of the model, results, and predictions are shown in online appendix E.1. We set the predicted start of the BCW as 10 years before the predicted retirement age, which we compute using the first age when an individual accumulates at least 50% likelihood of retirement. Then, we recenter the data around the predicted start of the BCW, extending the sample until 8 years after. This allows us to estimate the regression of equation (1), where event time (Years Window) is defined relative to the predicted start of the BCW. We allow event time -5 to include observations that are classified as earlier than 5 years before the start of the predicted BCW.[12]

The main identification assumption is that individuals in each event time group are on average identical, conditional on the set of controls. Thus, all differences in reported earnings across event time groups are assumed to come from the dynamic incentives of the pension system. Because event

[11] Sole-ownership firms with yearly revenue below a certain threshold are exempt from reporting detailed balance sheets to the internal revenue service, which makes it more difficult to monitor the correct reporting of worker earnings for this type of firm. They very rarely have ten employees or more. Reporting around ten employees working full time at the minimum wage would put the costs of a firm (and thus revenue, unless significant losses were reported) above the threshold for eligibility for these exemptions. This creates additional incentives for firms close to the threshold to under-report revenues and costs, including payroll costs.

[12] Alternatively, we can directly use the retirement sample, which has data on retirement and work histories, to conduct event studies around the actual start of the BCW. However, the characteristics of this sample are not well suited for this empirical analysis, due to its small sample size and because only early cohorts are represented in the sample. In online appendix E.2, we show that our results are robust in event studies around the observed retirement age for the retirement sample.

FIGURE 2.—EVENT STUDIES OF REPORTED EARNINGS TO SOCIAL SECURITY



The plots show estimates of the coefficients of event time dummies in event studies around each individual's predicted start of benefit calculation window (BCW). The spikes illustrate 95% and 90% confidence intervals based on robust standard errors clustered by individual. The outcome is reported earnings to social security, as a ratio of the sample average at event time −1. The sample includes (A) self-employed workers, (B) employees of firms with fewer than ten workers, or (C) employees of firms with ten or more workers, from age 45 until 8 years after the start of the BCW, reporting earnings for 6+ months, between April 1996 and April 2016. All regressions include month-by-year effects, dummies for seven categories of firm size when first observed and contemporaneously, dummies for two-digit industry codes when first observed, and interactions of year fixed effects with firm size categories and industry codes when first observed. Details on the prediction of the BCW are provided in online appendix E.

time is based on age, it is key to control for changes in earnings over time. The sample composition, which includes observations for 30 birth cohorts, observed between ages 45 and 57, between 1996 and 2016, provides sufficient variation in age conditional on observation year to allows us to separately identify age and time effects. A challenge to identification in this setting would come from changes in productivity or wages with experience, and thus age. The event study allows us to partially test this identifying restriction, by showing that there are no age effects before the start of the BCW. Therefore, if there were any confounding changes in earnings with age, they would have to happen only after the start of the BCW, which seems implausible.[13]

We can expect responses to gradually increase in the BCW, even in the analysis based on predicted retirement. There are several reasons why workers may increase earnings gradually as they progress into their expected BCW. If workers heavily discount the future, they may not find it worthwhile to invest in reporting higher earnings to increase retirement benefits until they are closer to retirement. Also, with uncertainty about future earnings, and given the existence of a minimum retirement benefit, workers may want to wait until being closer to retirement before making the investment of increasing reported earnings. Additionally, some self-employed workers face institutional restrictions on the frequency and amounts of changes in reported earnings that can delay their responses, as discussed in section II.

We conduct several specification and robustness checks, discussed in section VF. First, we estimate a more parsimonious specification of the event study model where we group event time effects in longer intervals, and show robustness to the set of controls included. Second, we estimate a specification with a linear trend in event time, where we model responses as a shift and trend break after the start of the BCW. We also conduct several additional robustness checks, including event studies with individual fixed effects, an analysis of sample attrition, dynamic difference-in-differences models using large-firm employees as a control group, a placebo test with employees of publicly owned firms, and an alternative measure of the start of the BCW.

[13] Implicitly, we also assume that an individual's birth cohort has no effect on earnings. It is a well-known problem that, without further assumptions, it is not possible to simultaneously identify age, time, and cohort effects, due to their multicollinearity. In our setting, age effects are the main object of interest, and time effects are of primary importance, because of the long period of time considered and the presence of business cycles. The concern for cohort effects is mitigated by the fact that we find no effects before the start of the benefit calculation window, to the extent that there is no reason to think that cohorts would start having an impact only after. We also show that the results are robust to a specification with individual fixed effects, which is identified with an additional normalization, as discussed in section VF.

B. Event Studies of Reported Earnings to Social Security

Our main approach to study reported earnings responses is to conduct the event studies described in equation (1) centered around the predicted start of each individual's BCW, that is, 10 years before predicted retirement age. In online appendix E.1, we present and discuss in detail the retirement age model, its estimation results and fit, and the retirement age predictions. In the main analysis sample, predicted retirement age ranges from 60 to 65, with the largest fractions of individuals predicted to retire at ages 61 or 62 (see online appendix figure E1).

Self-employed workers. We start the empirical analysis by focusing on self-employed workers, who have both the incentives to change their reported earnings and the opportunity to do so, because they can more easily manipulate reported earnings. The estimates and confidence intervals from this event study are shown in figure 2, panel A. The results indicate that there is no pretrend in the years leading to the start of the BCW. After its start, average reported earnings start increasing significantly, reaching a level around 6% higher than the year before the start of the BCW within the first 5 years.

Employed workers. The results for workers employed in firms with fewer than ten employees are shown in panel B of

figure 2. For this group of workers, we find similar increases in reported earnings after the start of the BCW, although less precisely estimated. The estimated responses are statistically significant after the second year into the BCW, and are close in magnitude to those of self-employed workers, reaching 6 percentage points by 6 years into the BCW. Reassuringly, our findings show that employees working at larger firms, shown in panel C, do not exhibit significant increases in reported earnings in the BCW.[14]

Event studies centered around age 50. In online appendix D.1, we show results for event studies where we center event time around age 50 for all workers, that is, 10 years before minimum retirement age. Although age 50 is a more noisy measure of the start of the BCW, the results are consistent with our main results. We find significant responses of self-employed workers soon after age 50, more imprecisely estimated responses for small-firm employees, and no evidence of a change in reported earnings for employees of larger firms. The results show that the increases in reported earnings occur gradually after the 50th birthday. For small-firm employees, estimated responses are larger after age 53, 10 years before the average retirement age in the sample. This suggests that some individuals expect to retire some years after the minimum retirement age, and thus start changing their behavior later than age 50.

While our main estimates suggest more rapid responses, we still find evidence of a ramp-up after the *predicted* start of the BCW. There are multiple possible explanations for this, including measurement error in predicted retirement age and institutional regulations that limit the speed of changes in reported earnings for self-employed workers. In addition, there can be behavioral reasons for delayed responses, such as time preferences (discounting of future retirement benefits when workers are further away from retirement), uncertainty about the likelihood and timing of retirement, or inattention combined with increased salience of retirement benefit incentives as workers approach retirement.

C. True Labor Supply Responses or Changes in Reporting Behavior?

So far, we have shown evidence of increased *reported* earnings during the benefit calculation window. To explore whether these changes in reported earnings are explained by changes in reporting behavior or actual increases in hours of work or total earnings, we estimate event studies around age 50, for survey-reported earnings and hours of work conditional on formal employment. The comparison of changes in earnings with age across the two sources—administrative records and household surveys—allows us to understand changes in earnings under-reporting as a response to the incentives of the pension system. In addition, by looking at hours of work as an outcome, we can directly test whether there are changes in labor supply after age 50.[15]

The results of this analysis are presented in figure 3. We do not find significant increases in hours of work or total earnings of the self-employed after age 50 (panel A). Although the estimates from household surveys are more imprecise, the point estimates suggest, if anything, a relative decline in self-employed earnings after age 50. In panels B and C, we present results for employees of small and large firms, respectively. We do not find any evidence of increases in hours of work or earnings of employees of small or large firms after they enter the BCW.

These results suggest that observed changes in reported earnings are most likely explained by changes in reporting behavior to the social security administration, rather than true earnings or hours of work. Moreover, the fact that we do not find evidence of changes in hours of work or total earnings with age is reassuring for the validity of our empirical strategy. If there were unobserved factors correlated with age that affect labor supply or earnings, we would expect them to also affect the estimated age effects in the household survey event studies. A key limitation is that the household survey data cannot be matched with the administrative data at the individual level to construct a direct measure of earnings under-reporting. Instead, we use the differences in reported earnings across the two sources and across age groups, by groups of workers and firm sizes, to provide suggestive evidence of under-reporting.

D. Heterogeneity among Self-Employed Workers

In this section, we explore the heterogeneity of responses across groups of self-employed workers that have different incentives and flexibility to respond. The results are presented in online appendix C.1, and summarized in online appendix figure C1.

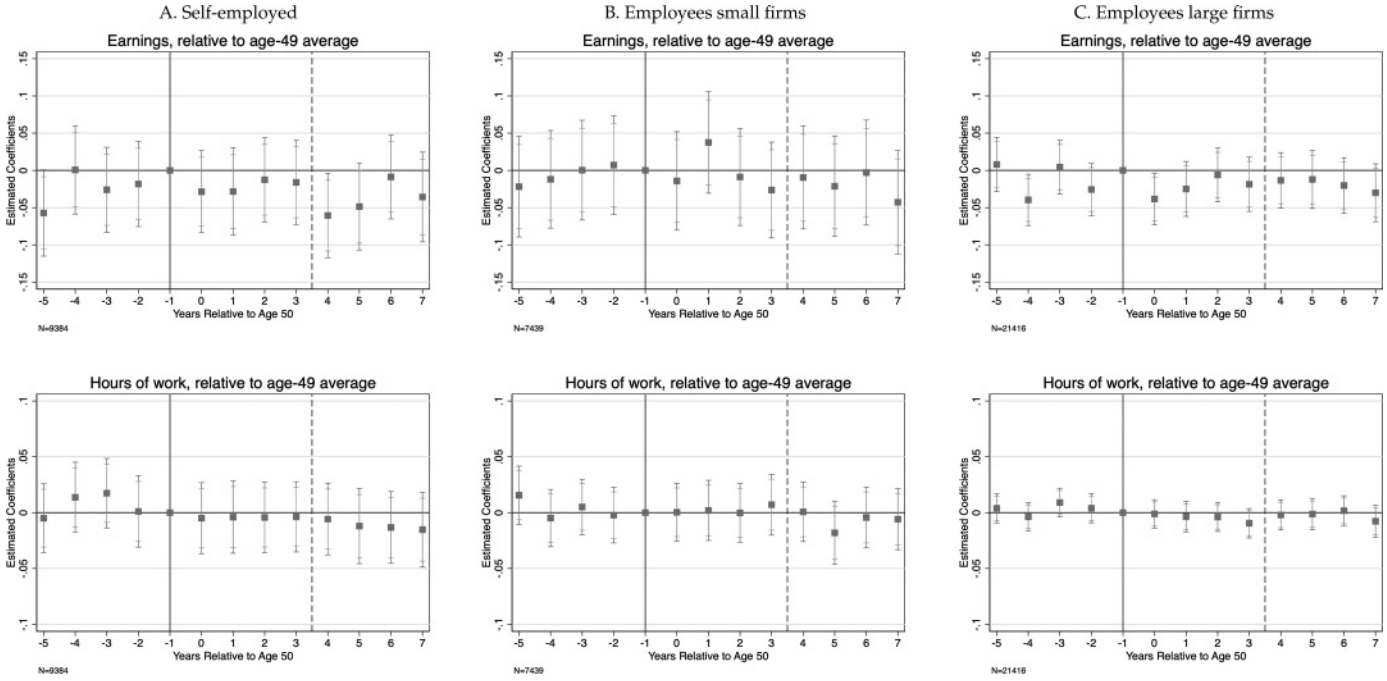
Minimum contribution base. Our model predicts "extensive margin" responses from self-employed workers who report the minimum contribution base (MCB), and "intensive margin" responses from workers who already report earnings above the MCB.[16] We conduct event studies to evaluate responses in these two margins. For the extensive margin, we estimate an event study where the outcome is an indicator for reporting the MCB. We find that the likelihood of reporting the MCB decreases by less than 2 percentage points in the BCW. For the intensive margin, we estimate changes in

[14]These results are consistent with the findings of Bergolo and Cruces (2014), who find reported earnings responses to formal-work benefits only for small-firm employees.

[15]In household surveys, there can also be earnings measurement error. Incentives to misreport are lower than in administrative data but the cost of misreporting is also lower (Hurst, Li, & Pugsley, 2014). However, there is no reason to change reporting behavior specifically after age 50.

[16]It should be noted that both of these margins are conditional on being registered as self-employed and reporting positive earnings. In section VF, we discuss the separate issue of gaps in earnings reporting.

FIGURE 3.—EVENT STUDIES OF EARNINGS AND HOURS OF WORK IN HOUSEHOLD SURVEYS



The plots show estimates of event time dummies in event studies around age 50. The spikes illustrate 95% and 90% confidence intervals based on robust standard errors. The outcomes are self-employed earnings (top) and weekly hours of work (bottom), relative to sample average at time −1. The sample includes workers ages 45 to 57 who are (A) self-employed, (B) employed in firms with fewer than ten workers, or (C) in firms with ten or more workers. Data from household surveys (ECH) April 1996 to December 2016. All regressions include month-by-year effects, dummies for four categories of firm size, dummies for two-digit industry codes, and interactions of year fixed effects with firm size categories and industry codes.

reported earnings on the subsample of self-employed workers who were reporting above the MCB at baseline. For this group, we find changes in reported earnings that are larger than for the full sample, and significantly larger than those of self-employed reporting exactly the MCB at baseline. These results suggest that a large part of the responses come from the intensive margin, that is, from workers who are already reporting above minimum earnings.

High formal self-employment attachment. We expect frequent self-employment reporters to benefit the most from adjusting their reporting earnings, because they are more likely to rely on self-employment as their main source of income and to meet the years of service required for retirement by their early sixties. To evaluate this, we study the subsample of self-employed workers with a density of contribution of at least 70% of the sample months. We find slightly larger point estimates of responses for this group than for the full sample of self-employed, but the difference is not statistically significant.

Earnings reporting regulations. As explained in section II, self-employed workers in transition cohorts are more limited in terms of the frequency and the magnitude of increases in reported earnings than cohorts in the new two-pillar regime. We find statistically significant differences in the responses of the two groups of cohorts, suggesting that the restrictions faced by transition cohorts significantly delay and reduce the magnitude of their responses. Self-employed workers in two-pillar cohorts have estimated responses that are about twice our baseline estimates for the full sample. We also study the heterogeneity of responses among self-employed workers with employees, whose reported earnings are tied to those of their employees. While this regulation is effective at increasing their baseline level of reported earnings, we do not find robust evidence that they are more or less responsive to the benefit calculation window.

Taken together, these results confirm that both incentives and opportunity to respond matter. We find stronger responses among self-employed workers who frequently report to social security, who have more clear incentives, and those with more flexibility to respond because they are not subject to regulations that limit changes in reported earnings.

E. Heterogeneity and Margins of Response among Small-Firm Employees

We conduct heterogeneity analyses to further explore the roles of firm size and formal employment attachment in shaping employed workers' responses to the benefit calculation window. We also explore margins of response, including increased earnings within the job and job changes, and changes in reported hours of work and implied hourly wages. The results are presented in detail in online appendix C.2.

Firm size. We begin by exploring the relevance of the ten-worker firm-size cutoff for separating worker responses.

TABLE 2.—REGRESSION ESTIMATES OF CHANGES IN REPORTED EARNINGS

	Self-employed			Small-firm employee			Large-firm employee		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
2–5 years pre-start BCW	−0.005	−0.005	−0.002	−0.021*	−0.002	0.005	0.023***	0.011*	0.011*
	(0.008)	(0.006)	(0.006)	(0.011)	(0.011)	(0.011)	(0.007)	(0.006)	(0.006)
0–3 years post-start BCW	0.017**	0.024***	0.028***	0.040***	0.029**	0.028**	0.003	0.007	0.008
	(0.008)	(0.006)	(0.006)	(0.013)	(0.012)	(0.012)	(0.007)	(0.006)	(0.006)
4+ years post-start BCW	0.054***	0.048***	0.062***	0.083***	0.055***	0.057***	0.009	0.021*	0.020*
	(0.015)	(0.011)	(0.011)	(0.021)	(0.019)	(0.019)	(0.012)	(0.011)	(0.011)
Observations	315,707	315,707	315,707	233,045	233,045	233,045	466,617	466,617	466,617
Time FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm size + Industry FE		Yes	Yes		Yes	Yes		Yes	Yes
Firm size + Industry # Year FE			Yes			Yes			Yes

Data from main analysis sample 1996–2016. Outcome: reported earnings over average at time −1. Samples include self-employed (columns 1–3), and employees of firms with <10 workers (columns 3–6), and of larger firms (columns 7–9), from age 45 until 8 years post predicted start of the BCW. Control variables: month-by-year effects (Time FE), seven categories of firm size when first observed and contemporaneous and two-digit industry codes when first observed (Firm size + Industry FE), and interactions of year fixed effects with firm size and industry when first observed (Firm size + Industry # Year FE). Robust standard errors clustered by individual in parentheses. $^{*}p < 0.1$, $^{**}p < 0.05$, and $^{***}p < 0.01$.

We estimate separate regressions for four firm-size categories: fewer than 5, 5 to 9, 10 to 49, and 50 or more workers. We find similar estimated responses for workers in the first two groups, while the estimates for the two groups of larger firms are close to zero. Although the confidence intervals overlap, the estimates suggest economically significant responses only in firms with fewer than 10 employees.

High employment attachment. We expect workers with stronger attachment to formal employment to have more incentives and opportunity to increase reported earnings. To assess this, we study the subsample of employees of small firms with a density of contribution of at least 70% and at least 1 year of tenure at the job at age 49. We find statistically significant changes in reported earnings after the start of the BCW, which are about twice as large as our main estimates for the full sample of small-firm workers.

Responses conditional on staying at the job versus job changes. If responses are enabled by long-term agreements between employer and employee, we expect stronger responses among workers who stay at the same firm. To assess this, we further limit the sample of small-firm workers with strong attachment to observations at the same firm where the worker was employed at age 49. The estimates are even larger, suggesting that workers who stay at the same firm are more likely to see their reported earnings increased. Alternatively, workers could achieve increases in reported earnings through job changes. We do not find any evidence of increased job switching after the start of the BCW.

Reported hours of work and implied wage per hour. Changes in reported earnings can be the result of changes in reported hours or earnings conditional on work hours. We explore these two margins by estimating event studies for reported hours by month and implied earnings per hour. Although there is little variation in reported work hours, which are mostly reported at typical full-time work schedules, we find some evidence of increases in reported hours of work in the BCW for small-firm workers. However, most of the estimated effects on reported earnings for small-firm workers can be accounted for by increases in implied hourly wages.

Altogether, the results are consistent with the prediction that stable employment relationships in small firms facilitate workers' ability to respond to the incentives of the pension system. Stable employment may favor long-term agreements between employer and employee, both in terms of true earnings or reporting behavior. The findings of only small changes in reported hours of work and no changes in work hours or earnings in household surveys suggest reporting behavior as the more likely margin of response. However, because of data limitations, we cannot completely rule out other possible response mechanisms, such as, for example, late career bonuses or job promotions.

F. Robustness

In table 2, we present a more parsimonious specification of the event study model where we group event time effects in longer intervals, again leaving event time −1 as the omitted category:

$$Y_{it} = \alpha \text{Pre BCW}_{it}^{2-5} + \beta \text{Post BCW}_{it}^{0-3} + \lambda \text{Post BCW}_{it}^{4-7} + \delta_t + \gamma X_{it} + \varepsilon_{it}, \quad (2)$$

where $\text{Pre BCW}_{it}^{2-5}$ is an indicator for observations 2–5 years before the start of the BCW, and $\text{Post BCW}_{it}^{0-3}$ and $\text{Post BCW}_{it}^{4-7}$ are indicators for observations in years 0–3 and 4–7 after the start of the BCW, respectively. We show robustness to the set of controls included in X_{it}, which always includes a dummy for observations more than 5 years before the predicted BCW. We start with only time (month-year) effects. Then, we add indicators for intervals of firm size and two-digit industry codes. Finally, we add a full set of interactions between calendar year and firm size and industry dummies, which is the same set of controls we use in the fully dynamic event study.

Including the full set of controls reduces the estimates for the pre-BCW coefficients and improves the precision of the

estimates. The results of this preferred specification indicate that self-employed reported earnings increase by 2.8% in the first 4 years of the BCW, and by 6.2% in the following 4 years, relative to time −1 (significant at 1%). Assuming reported earnings are stable in the last 2 years of the BCW, these estimates imply an average reported-earnings increase of 4.8% in the 10-year BCW. The estimated changes in reported earnings for small-firm employees are similar to those for self-employed workers.

In online appendix D.2, we present estimates of a specification with a linear trend, where responses are modeled with a shift and a trend break in the BCW, which we allow to be quadratic. The estimates indicate that, after the start of the BCW, self-employed reported earnings increase significantly and at a decreasing rate. The presence of a trend break rather than a shift at exactly the start of the BCW is consistent with the results of the fully dynamic event study. For small-firm employees, the estimated trend break is also positive, although not statistically significant.

In online appendix D.3, we estimate event studies that include individual fixed effects. This allows us to control for permanent individual unobserved heterogeneity, including cohort effects. To allow for identification of the event-time effects conditional on individual and time effects, we drop an additional pre-event time effect (Borusyak, Jaravel, & Spiess, 2021). We find that our results are robust to the inclusion of individual fixed effects.

A potential concern is the possibility of differential sample attrition across groups of workers. The fixed-effects model mitigates this concern by controlling for unobserved individual heterogeneity. In online appendix D.4, we directly examine whether there is a pattern of differential sample attrition across groups. We balance the samples by adding zeros in periods of nonreporting, and estimate difference-in-differences event studies where the outcome is an indicator for being in sample, using large-firm employees as a control group. The results show no significant differential attrition among the self-employed or small-firm employees, relative to large-firm employees.

In online appendix D.5, we estimate effects on reported earnings using large-firm workers as a control group in dynamic difference-in-differences analyses for self-employed and small-firm workers. This specification identifies the differential responses of these two groups to the BCW relative to those of employees of larger firms, which are less prone to earnings under-reporting. It allows for the presence of unobserved changes in reported earnings with age, under the assumption that these changes are common to all workers. The results confirm our findings, with only small changes in the estimated responses.

In online appendix D.6, we conduct a placebo test on workers of publicly owned firms. These workers are excluded from our main sample because they are public-sector employees. Although they are also affected by a 10-year BCW, we should not find responses among this group, for two reasons. First, there is no under-reporting of earnings in the public sector. Second, publicly owned firms are very large firms subject to several layers of administrative monitoring, making responses challenging in terms of increased compensation. Reassuringly, we do not observe any significant changes in reported earnings in the predicted BCW for this group.

In online appendix D.7, we present an event study for self-employed workers where the dependent variable is reported earnings as a fraction of the MCB. This allows us to test the sensitivity of the estimates to changes in the MCB, and at the same time provides an alternative way of expressing the magnitude of the effects. The results follow the same patterns as our main results. After 4 years into the BCW, self-employed reported earnings increase by 7.3% of the MCB relative to the year before the start of the BCW.

In online appendix D.8, we conduct additional robustness checks on the household survey analysis. First, we estimate semiparametric and parametric age-trend models, similar to those of table 2 and online appendix D.2, with different sets of controls. Second, we explore the robustness to using as measure of earnings the net earnings directly reported in the survey, before imputing social security contributions to obtain gross earnings. Across all robustness checks, we find no evidence of increases in earnings or hours of work for any group of workers. We also conduct an analysis on the self-reported information about earnings under-reporting to social security, available for employed workers since 2006. We find suggestive evidence of decreased under-reporting for small-firm workers, and not for large-firm employees. This is consistent with the conclusions we extract from comparing our analysis of the administrative and survey data, which suggest that responses occur along the margin of changes in under-reporting behavior.

What happens at a firm when a worker enters the benefit calculation window? One of the challenges for identifying employee responses to the incentives of the pension system is the possibility that the observed changes in reported earnings could be related to changes at the firm level, such as increases in firm productivity or labor demand. In online appendix D.9, we study what happens at the firm around the time when one of its workers enters the BCW, using matched employer-employee social security records from the *firms sample*. We identify the first worker who turns age 50 in a firm as the firm's *reference worker*.[17] First, we reproduce the results obtained in our main sample for the reference worker. Then, we study other firm-level outcomes. We do not find any evidence of a pretrend in employment or the wage bill before the reference worker's 50th birthday. After the first worker turns 50, the total wage bill increases consistently with the increase in reported earnings for the reference worker. We do not find any evidence of changes in the

[17] We use age 50 as the start of the BCW, because in the firm data set we do not have each employee's work history, which is necessary to predict retirement age.

number of workers employed by the firm. These results suggest that the change in reported earnings for the reference worker is not due to changes in the firm's labor demand. Additionally, if firms absorb part of the cost from increasing reported earnings, this is at least not an unexpected shock that moves the firm from its growth trend.

Predicted retirement age. In online appendix E.2, we also show that the estimates of event studies of reported earnings around the predicted start of the BCW are robust to regressions where we weight observations by each individual's predicted probability of retiring at age 60, and to the use of an alternative measure of predicted retirement age, using the expectation obtained from the predicted probabilities of retirement by age.

VI. Back-of-the-Envelope Calculation of Costs

To provide a sense of the economic importance of the responses to the dynamic incentives of the benefit calculation window, we present "back-of-the-envelope" calculations of the implied costs to the pension system. We propose a simple model to compute these costs, which we present in detail in online appendix F.1. We focus on the self-employed, because this is the group with the most robust evidence of changes in reported earnings in the BCW. To the extent that at least some small-firm employees also respond, these calculations provide a lower bound to the actual implied costs.

We provide estimates under three scenarios of what the counterfactual reported earnings dynamics would be in the absence of responses. Without responses to the BCW, there would be no differentiation of reported earnings between the period before (period 1) and the period after the start of the BCW (period 2). In scenario 1, we assume that, in the absence of responses, self-employed workers would continue reporting the same earnings that we observe them reporting before the BCW. This implies that responses take the form of increased reported earnings in the BCW, relative to what they would otherwise report. The cost to the social security system comes from an excess expenditure in retirement benefits, partially mitigated by the increase in payroll tax revenues during the BCW. In scenario 2, we assume instead that reported earnings would be as those reported in the BCW. This means that responses take the form of a decrease in reported earnings in period 1, and the cost to the social security administration comes exclusively through a loss of revenues during this period. In scenario 3, we consider an intermediate case in which reported earnings, in absence of responses, would be the average between those observed before and after the start of the BCW. In this case, the cost to the social security administration comes from two mechanisms: an increase in expenditures during the retirement period, and a decrease in revenues during period 1, partially compensated by an increase in revenues during period 2.

We calibrate the model considering the characteristics of the institutional setting. In our main estimates, we assume that the average self-employed worker retires at age 63, with 33 years of service. The BCW (period 2) is the 10-year period prior to retirement. The previous 23 years of service happen continuously in the years preceding the start of the BCW. The duration of period 3 is the male life expectancy at the age of retirement. The payroll tax rate is 22.625%. We assume a real interest rate of 2%, which is about the average of the last 30 years. For the replacement rate, we use the statutory replacement rates for the retirement age and years of service considered (de Melo et al., 2019). For the computation of factual and counterfactual reported earnings, we use as benchmark reported earnings in period 2, the sample average self-employed reported earnings at ages 54 to 57. We compute the increase in reported earnings in the BCW using a weighted average of our estimated changes in reported earnings in the BCW, relative to the year before the start of the BCW (4.8 percentage points). To quantify the costs in millions of U.S. dollars of 2019, we calculate the cost per self-employed individual in dollars, annualize it, and multiply it by the number of self-employed retirees in the main pension system. Finally, we also present the cost as a percentage of Uruguay's 2019 GDP.

Under scenario 1, we estimate that net social security expenditures increase by 3.3% per self-employed worker, which represents US$103 million per year or 0.18% of the GDP. In the second scenario, the estimated net cost is larger, with an increase of 4.2% in the cost per self-employed retiree, equivalent to US$118 million and 0.22% of the GDP. Scenario 3 represents an intermediate case between the first two scenarios, with an estimated excess cost of 0.20% of the GDP.

In online appendix F.2, we show additional estimates under various alternative parameterizations. We vary retirement age from 60 to 65, and consider a counterfactual with gaps in the contribution history. The estimated costs in all scenarios vary little with retirement age and implied years of service. Then, we consider our baseline parameterization under alternative interest rates, of 0% and 4%. The direction of the changes of estimated costs with the interest rate depends on the scenario considered, given the different importance of future expenditures and past contributions in each scenario. Overall, the results are robust to this wide range of parameterizations, with estimated costs that vary from 0.13% to 0.35% of the GDP.

VII. Final Remarks

Many pension systems in developed and developing countries take into account reported earnings in a small set of final years of work in the computation of pension benefits. This decision introduces dynamic incentives for individuals to shift reported earnings toward the final years of their careers to increase expected pension benefits. The presence of responses to these dynamic incentives is not trivial, since workers may have optimization frictions and may face high costs of changing their reporting behavior in their

final years of work. In this paper, we use matched social security administrative records from Uruguay, supplemented with household surveys, to study the earnings-reporting behavior of self-employed and employed workers in an event-study framework. To the best of our knowledge, this is the first paper to show that self-employed workers and some private-sector employees respond to the dynamic incentives of the retirement system by increasing reported earnings in the benefit calculation window. Furthermore, we provide evidence that suggests that the main mechanism that explains these changes in reported earnings is a change in under-reporting behavior.

Self-employed workers are a group that is particularly likely to respond to these incentives, because their earnings reporting is not mediated by employers with whom to bargain and they are more prone to under-reporting. Our findings suggest that also some employees of small firms, mediated by their employers, can change reported earnings in response to dynamic incentives. We find evidence that these responses are more likely in long-term employment relationships, and that they do not represent an unexpected shock for firms. This is consistent with the possibility that responses operate through implicit or explicit long-term dynamic contracts on reported earnings. It remains for future work to theoretically explore and empirically evaluate more directly the bargaining process between workers and firms on compensation and earnings reporting to social security in the response to dynamic incentives from taxes and transfers.

Pension design should consider the excess cost created by responses to the dynamic incentives induced by a short benefit calculation window. Our estimations imply that these responses impose substantial costs to the pension system, increasing net expenditures by about 0.2% of the GDP, which is roughly equivalent to government expenditures in early child care. However, this cost estimation is not a welfare assessment. In such an assessment, utility gains for workers who respond to these incentives would have to be weighed against utility costs for other taxpayers. It remains for further research to explore the potential of different policy instruments to address this issue. Extending the window of benefit calculation could mitigate this cost by removing the dynamic incentives, improving fairness. However, it is possible that the extension of the window would not achieve the goal of increasing reported earnings. Moreover, it could further reduce future retirement benefits of the workers who respond to these incentives, who already have low expected pension benefits. Alternatively, improvements in reporting regulations and enforcement could help improve the levels of retirement contribution and benefits, especially among the self-employed, but can also increase the cost of formal self-employment.

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