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The Impact of Agricultural Extension for Improved Management Practices: an Evaluation of the Uruguayan Livestock Program

Conner Mullally and Alessandro Maffioli[1]

Abstract:

Management is an important input into agricultural production, as it a determinant of the uptake and proper implementation of productive technologies and practices. While there is a large literature on evaluations of extension programs meant to improve management practices in agricultural development, there is no consensus on the extension modalities that are most effective. This paper adds to the literature on extension interventions by evaluating the Uruguayan Livestock Program (ULP), a publicly funded, privately delivered extension program designed to improve management practices among cattle breeders. Using an eight year panel constructed by combining data from the Uruguayan livestock traceability system with a registry of ULP participants, we place bounds on the impact of the program on production and sales of calves by ULP beneficiaries using inverse probability weights estimated using propensity scores for selection into the ULP and selection into the dataset. Results show that the ULP increased calf production by between 11.36 and 15.3 calves on average in 2009 and 2010 and increased calf sales minus purchases by 4.35 on average over the same time span. Internal rates of return suggest these are moderately sized economic impacts. We examine the aspects of the ULP's design that might account for its positive but modest effects.

JEL codes: M11, O13, 022, Q12, Q16

Keywords: agricultural extension, agricultural production, impact evaluation, livestock

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1. Introduction

Economic theory offers several explanations for the persistence of poor management practices in developing countries. For example, firms that produce at a lower average cost than their competitors as a result of adopting best management practices may not be able to expand because of failures in input markets or liquidity constraints. At the same time, market failures may affect the flow of information on good management practices. For example, transaction costs faced by information providers may combine with uncertainty over the value of information among potential clients, resulting in a lower quantity of information on management practices transacted than would be the case if markets functioned perfectly (Anderson & Feder, 2004).

Given the above it is no surprise that there is a strong interest in development economics in the impacts of interventions designed to improve management practices on firm performance. In agricultural development, there is a rich literature on agricultural extension, and agricultural extension programs generally focus on farmer management to some degree. This paper adds to the literature on agricultural extension in general and management interventions in particular by evaluating the Uruguayan Livestock Program (ULP), a program consisting of publicly funded, privately delivered extension services for cattle breeders.

We estimate the average impact of the ULP on the number of calf births and net sales of calves by beneficiaries in 2009 and 2010, focusing on ULP beneficiaries entering the program in 2008 who own at least some land. We estimate impacts using inverse probability weighting; the weights are estimated using all pre-program years observed in an eight year panel dataset constructed from two different sources of administrative information from the Uruguayan Ministry of Livestock, Agriculture, and Fisheries (MGAP). The weights are constructed using the propensity score for participation in the ULP and the propensity score for selection into the dataset in the year for which impacts are estimated; in the administrative dataset used for the analysis, cattle producers are observed in some but not all years, and we assume that any potential biases arising from selection into the dataset can be controlled for using observed information.

Our estimation method point-identifies the average treatment effect on the treated (ATT) for net calf sales (sales minus purchases) and identifies upper and lower bounds on the ATT for calf births; the ATT for calf births is partially identified rather than point identified because of the presence of missing data. Our results indicate that the ULP increased calf births among program beneficiaries by between 5.1 and 7.58 in 2009 and between 5.46 and 6.78 in 2010 on average, while the average impact on net calf sales among beneficiaries was 1.6 in 2009 and 2.83 in 2010; all but the impact on net calf sales in 2009 are statistically significant at conventional levels. The magnitude of these effects is moderate, as the market price of a calf in Uruguay is around $150 and the average subsidy per beneficiary was $3,259. We subject our method to robustness checks by estimating "placebo" ATT parameters for calf births and net calf sales in years prior to the start of the program. The results of these robustness checks suggest that estimated impacts for 2009 are accurate while those for 2010 may be biased. Finally, we also calculate a rough estimate of the internal rate of return (IRR) of the program; the results are sensitive to the time horizon over which benefits are assumed to last and the outcome used to value program benefits.

In what follows, we present a brief review of the literature on impact evaluations in agricultural extension. We then describe the motivation and design behind the ULP program. We then present the empirical method used to point identify the ATT on calf sales and partially identify the ATT on calf births, which is followed by our results and robustness checks. We conclude with a brief summary and directions for future research.

2. The impacts of agricultural extension

This paper adds to the rich literature in agricultural economics on the evaluation of extension programs. Evenson (2001) reviews impact evaluations of extension projects produced through the 1990s, and finds a median estimated internal rate of return (IRR) for farm-level impact studies of 80%. Many of the studies cited by Evenson measure the average *ceteris paribus* impact of extension by using cross-sectional data to estimate a structural production function, requiring proper specification of the production function and unbiased estimates of its parameters. In contrast, most of the more recent literature on agricultural extension combines panel data with a program evaluation approach, i.e., comparing average outcomes in the treatment group (producers receiving extension services) and a control group (producer who do not receive extension services) in order to estimate an average treatment effect (the average impact on the program on a population of interest, usually the treatment group). Identification of extension impacts under the program evaluation approach requires identification of a single parameter in order to generate unbiased estimates of extension impact, i.e., the counterfactual average outcome in the treatment group in the absence of extension. Estimated impacts will also include effects resulting from changes in other aspects of the production function, whereas *ceteris paribus* estimates would net out any such effects.

Reviews of extension impact evaluations that include an explicit counterfactual include IEG (2011) and IDB (2010). Extension modalities evaluated include farmer field schools set up to teach producers about Integrated Pest Management (IPM), visits to individual farms by extension agents, publicly funded and privately delivered extension advice geared towards a single commodity (as in this paper), and provision of market information through cell phones. The most commonly used indicators capture impacts on knowledge accumulation (e.g., answering questions about IPM correctly) or production and earnings (e.g., yields, profits, revenue). While virtually no extension interventions are shown to have negative impacts in these literature reviews, it is just as likely that a study fails to reject the null hypothesis of no effect as it is that the estimated average impact is both positive and statistically significant; no single extension modality has a large enough body of evidence to safely conclude that it is the most effective. Studies of extension using the program evaluation approach do not typically calculate an IRR.

Several evaluations of extension programs have been published subsequent to the reviews cited above. These include Maffioli et al. (2013), who find that extension services have positive impacts on technology adoption but no significant effect on yields for small and medium-sized fruit producers in Uruguay; Cunguara and Moder (2011) in Mozambique, who find that extension services raise incomes among beneficiaries by an average of 12%; Bellemare (2010), whose estimates of the elasticity of yields with respect to extension visits fall between 1.3 and 1.7 for contract farmers in Madagascar; Lapple, Hennessy, and Newman (2013), who find that participatory extension has positive average impacts on profits for Irish dairy farmers; Olaganju and Adesiji (2011), who are unable to reject the null hypothesis of no effect of extension cocoyam yields in Nigeria; Davis et al. (2012), who find large and significant impacts of farmer field school participation on production and income in Kenya and Tanzania but not in Uganda; and Goodhue, Klonsky, and Mohapatra (2010), who find that extension focused on pesticide use results in a moderate but significant average reduction in the probability and intensity of pesticide use among California almond producers. As in the case of the earlier literature, extension does not appear to have negative impacts, although effects are not always statistically significant; whether this is because of small sample sizes or because the true underlying impacts are small is unclear.

To the extent that studies using an explicit counterfactual yield more credible results, the quality of the evidence on extension impacts has improved. However, more evidence is needed before a consensus can be reached on the effectiveness of different extension modalities. Our paper benefits from the availability of a large number of observations, addressing concerns over statistical power. We also exploit the presence of multiple years of pre-program data by estimating “placebo” treatment effects, i.e., program impacts in years prior to the start other ULP. Rejecting the null hypothesis of no effect in a pre-program year would suggest that our empirical approach does not eliminate selection bias.

3. The Uruguayan Livestock Program: context and program design

Beef production is the primary agricultural activity in Uruguay and the top exported commodity with total exports of around $1.3 million annually (MGAP, 2012). Uruguayan beef cattle are almost exclusively fed on pasture from weaning until slaughter, are free of hormones by law, and produced primarily for export. Generally speaking the supply chain consists of breeders, fattening for slaughter, and “complete cycle” producers who engage in breeding and fattening; a small number of producers specialize in feeding cattle from weaning until reproductive age or until they are ready to be fed more intensively prior to slaughter. Breeding is by far the dominant activity in the supply chain, as around 15,000 of the 25,000 cattle ranches operating in a given year are either dedicated to breeding or complete cycle production. Nearly all producers engage in fattening for slaughter to some extent, but producers who specialize in cattle fattening very rarely engage in cattle breeding (Saravia, César, Montes, Taranto, & Pereira, 2011).

Management styles for grass-fed cattle breeding can be thought of as lying along a continuum. At one extreme is "poor management," characterized by a lack of record keeping, year-round breeding, minimal effort to maintain the reproductive health of cows and bulls, maintaining animals in a single fenced area, possibly with a small area of pasture reserved to prevent animal starvation, and weaning calves just prior to sale. At the other extreme of "good management," where detailed economic and production records are kept and the breeding season is organized around pasture availability; cows are exposed to bulls just as winter is beginning so that the high nutritional needs of lactation occur simultaneously with maximum pasture availability in the spring. In addition, animals are separated by gender and body condition into distinct fenced lots so that nutritional needs and breeding are more easily managed. Weaning is timed so as to allow breeding cows to recover their nutritional status in time for the next breeding cycle. The reproductive health of bulls is verified prior to breeding, and after exposure cows are pregnancy tested; cows that fail to become pregnant are either culled or fattened for slaughter. Lastly, good managers practice sustainable pasture management by not overloading land with livestock. Most of these practices are included in the package of management techniques promoted by MGAP to cattle breeders throughout the country (Saravia, César, Montes, Taranto, & Pereira, 2011).

The result of good management practices should be a ratio of calves weaned to breeding cows that is reasonably close to 100%; in Uruguay this figure is usually around 60%, implying that in any given year around 40% of the breeding stock is using resources without producing offspring (MGAP, 2012). This suggests that management practices among cattle breeders in Uruguay are in general quite poor. This conclusion was confirmed by the National Livestock Survey (NLS), a nationally representative survey of livestock producers with at least 100 hectares of land carried out in 2001 and 2003. Table 1 below summarizes data from the 2003 NLS on implementation of key management techniques by size of landholdings.

Table 1. Management practices by size of landholdings

Management indicators (1 = Yes, 0 = No)	Means (1) 100 - 500 ha	(2) 500 - 1,250 ha	(3) >1,250 ha	Differences in means (3)-(2)	(2)-(1)	(3)-(1)
Economic registry	0.527 [0.051]	0.716 [0.050]	0.884 [0.028]	0.168 [0.057]**	0.189 [0.072]**	0.527 [0.051]**
Production registry	0.456 [0.051]	0.586 [0.058]	0.755 [0.035]	0.169 [0.068]*	0.13 [0.077]	0.299 [0.061]**
Uses public or private extension regularly	0.147 [0.035]	0.292 [0.055]	0.499 [0.035]	0.207 [0.065]**	0.145 [0.065]*	0.352 [0.050]**
Seasonal breeding	0.731 [0.044]	0.936 [0.025]	0.932 [0.017]	-0.004 [0.030]	0.206 [0.050]**	0.202 [0.047]**
Separates animals by gender and fertility	0.136 [0.034]	0.197 [0.044]	0.353 [0.032]	0.156 [0.054]**	0.061 [0.056]	0.218 [0.047]**
Classification by body condition	0.543 [0.051]	0.591 [0.058]	0.607 [0.035]	0.016 [0.068]	0.048 [0.077]	0.063 [0.062]
Early/temporary weaning	0.576 [0.049]	0.573 [0.058]	0.577 [0.035]	0.003 [0.068]	-0.003 [0.076]	0.001 [0.060]
Pregnancy testing	0.235 [0.042]	0.506 [0.059]	0.691 [0.034]	0.185 [0.068]**	0.271 [0.072]**	0.456 [0.054]**
Checkups for bulls	0.359 [0.048]	0.587 [0.058]	0.757 [0.032]	0.17 [0.067]*	0.228 [0.076]**	0.398 [0.058]**
Observations	117	103	472	575	220	589

Source: 2003 Uruguayan Livestock Survey; standard errors in brackets, *p<0.10, **p<0.05, ***p<0.01

The first three columns of table 1 show means of binary indicators of management practices for farmers with between 100 and 500 hectares (the latter is the legal cutoff for status as a “family farm” in Uruguay), 1,250 hectares (the eligibility cutoff for the ULP), and beyond. The next three columns show differences in management indicators across the different landholding categories. There is heterogeneity in the rate of adoption of the various practices, both across practices and landholdings. With the exception of practicing early or temporary weaning and classification by body condition, all management practices are strongly correlated with the size of landholdings. In addition, the regular use of extension is much more common among larger producers, with a difference of 35 percentage points in the usage rate of farmers in the largest land cohort versus that of farmers with less than 500 hectares. In general, large producers employ good management practices, but producers with smaller landholdings do not. Based on the 2001 NLS, producers with fewer than 1,250 hectares manage 44% of the breeding stock and 40% of the cattle stock overall. This suggests that in addition to possible concerns over equity, poor management practices among smaller producers may have implications for the sector as a whole in the form of high prices for calves.

In response to apparent shortcomings in management practices among livestock producers, MGAP and the Inter-American Development Bank created the Uruguayan Livestock Program. The ULP began in 2001 with a pilot phase that focused exclusively on producer business plans. The pilot program was evaluated by Lopez and Maffioli (2008) using difference-in-differences with propensity score matching as applied to the NLS panel data set matched to program monitoring data; results showed impacts of 18 and 25 percentage points on the percentage of beneficiaries keeping records of physical and economic events, respectively, but no significant productivity effects.

In 2007, a new version of the ULP was implemented that ran through 2010. The new ULP included several new components (value chain development, improvement of the livestock traceability system, incorporation of forestry into livestock operations, and promotion of producer groups) as well as a revised version of the cattle breeder business plan component. In order to be eligible for the program, producers had to operate fewer than 1,250 hectares of average quality as measured by the CONEAT index.[2] Cattle breeders wanting to submit a business plan first had to pick one of 49 private extension agents from a list maintained by ULP managers. Extension agents were trained on how to put together business plans and submit them to MGAP before being eligible to work with beneficiaries. A complete business plan included a written proposal describing how the beneficiary would overcome current obstacles through carrying out the business plan, a 30 month calendar of activities and their costs, two and five year projections of sales, purchases, and holding of livestock with and without the business plan, the nature and timing of goals to be reached (including the goals that would trigger subsidy payments), and a detailed inventory of current livestock mirroring the data submitted annually to the national livestock traceability system.

In exchange for their participation, extension agents received 10% of the subsidy given to each producer, a fee of $100 for assisting with the development of a business plan, and a fee of $75 for each of two reports detailing the progress of individual beneficiaries towards reaching business plan goals. The maximum subsidy was $4,000 (down from $7,000 during the pilot) or 50% of the total cost of the plan, while the average total payment was $3,259. Business plans were to be selected on the basis of net present value, the likelihood of reaching stated goals, and possibilities for dissemination. Plans were pre-screened by the contracted extension agents with final approval by ULP managers. A total of 28.2% of business plans were audited during development or execution; one extension agent was dismissed because of inconsistencies between audit results and filed reports (Rearte, 2011). The details of finalized business plans for all ULP cohorts are summarized in Table 2.

[2] The CONEAT index assigns numbers to all agricultural land parcels in Uruguay based on soil productivity. The average value is 100. Parcels with scores between 80 and 120 are suitable for livestock production. More information can be found at http://www.prenader.gub.uy/coneat/doc/doc_coneat.htm.

Table 2. Details of completed business plans

Business plan goals	Count	Percent of all goals	Percent completed
Reproductive performance	**2,441**	**33%**	**87%**
Increase weaning % (calves weaned/breeding cows)	703	9%	81%
Increase pregnancy % among cows	593	8%	95%
Increase pregnancy % among ewes	326	4%	83%
Increase the size of the cattle breeding stock	304	4%	89%
Increase number of calves per hectare	129	2%	81%
Breeding management	**1,682**	**22%**	**94%**
Pregnancy testing	363	5%	91%
Early weaning	332	4%	98%
Veterinary checkups for bulls	289	4%	93%
Seasonal cattle breeding	268	4%	96%
Temporary weaning	167	2%	93%
Pasture management	**1,569**	**21%**	**89%**
Additional shade	587	8%	86%
Improved pasture	359	5%	89%
Organizing herd in separate areas by age and gender	150	2%	92%
Planting grain feed	148	2%	95%
Fertilization	99	1%	94%
Business management practices	**820**	**11%**	**98%**
Technical assessment of operation	420	6%	98%
Forming informal producer groups	191	3%	100%
Receiving additional technical training	154	2%	95%
Bookkeeping	55	1%	94%
Infrastructure	**679**	**9%**	**94%**
New electric fencing	169	2%	92%
New non-electric fencing	165	2%	95%
Improved water sources	104	1%	93%
Division of parcels into lots	96	1%	96%
Installations, e.g. loading and handling facilities	71	1%	97%
Other	**243**	**3%**	**90%**
Total	**7,481**	**100%**	**90%**

Source: MGAP

Table 2 organizes business plan goals into broad categories and then lists the most frequently cited activities and sub-goals associated with each goal. The vast majority of business plan goals are related to livestock reproduction, particularly cattle reproduction. Changes in pregnancy rates among the breeding stock, along with expansion of the breeding stock, should translate into more calf births, while increasing the ratio of calves weaned to cows inseminated should result in higher calf sales if the breeding stock is not reduced.

A total of 1,322 individual business plans were approved across four program cohorts, and 1,058 plans were completed. Beneficiaries were spread out over four cohorts. The program was largely the same for each of the four cohorts, with the only exception being that the first program cohort did not receive part of its subsidy payment up front. Cohorts 2, 3, and 4 received 25% of the total payment shortly after starting the program, 25% after achieving intermediate goals, and 50% upon completion. Cohort 1 received half of the total payment after reaching intermediate goals and the other half upon completion of their business plans, as in the pilot project.

4. Data

The data using in our analysis come from two separate components of the livestock traceability system operated by MGAP. The first of these is managed by the Livestock Control Division of MGAP, known by its Spanish acronym DICOSE. Under the DICOSE system, producers and their owned plots of land are identified by numeric codes. Producers are obligated to make a “declaration” each July in which they list the number of each type of animal owned on June 30 of that same year, tenancy and use details for each plot of land used for livestock production at the time of the declaration, and additional details such as consumption of owned livestock, animal births, and animal mortality rates.

All producers who own more than 10 cows or 50 sheep or have them in their custody for any other reason on June 30 (renting out grazing land, import and export) are obligated to make a declaration (MGAP, 2001). Submitting a declaration carries no cost beyond filling out paperwork and submitting forms at the local police station or MGAP offices. Producers are audited at random to monitor compliance, and non-compliance runs the risk of fines that escalate depending on the size of reporting errors and suspension of a producer’s DICOSE code, which is required for making sales and purchases. That being said, non-compliance does occur at a non-negligible scale. For example, 7,373 of around 52,000 DICOSE numbers were suspended in 2012 for non-compliance (Agromeat, 2012). DICOSE data in any given year may therefore not include livestock portfolios for all producers obligated to make a declaration.

In addition to annual declarations of land and livestock holdings, all movements of livestock, whether for the purpose of a transaction or without a change in ownership, require involved parties to file “Property and Transit Guide” forms containing the DICOSE codes of the individuals sending and receiving the animals, the DICOSE codes of the land parcels or facilities from which the animals are leaving and where they will arrive, mode of transport, dates, and other information. Since 2006 producers have also been obligated to identify each animal with an electronic ear tag containing relevant DICOSE codes and other information and register each animal under a unique code with MGAP; these animal codes must also be included with the information filed as part of any transaction (MGAP, 2006). Since the introduction of electronic tags, all transactions have had to be authorized by a representative of MGAP (which could be a producer involved in the transaction, if properly credentialed) and the data for all animals involved sent electronically to the traceability data center prior to movement of any animals. Transaction information is managed by the National System of Livestock Information, known by its Spanish initials SNIG. As in the case of DICOSE, non-compliance with SNIG also carries fines that escalate depending on the size of differences between animal quantities transported and what is reported on forms.

For this study, producer DICOSE codes were used by MGAP to construct a panel of cattle breeders, feeders, and “complete cycle” producers who own land, i.e., producers only reporting zero or missing values for owned land were filtered out of the database by MGAP. The data include the quantity of animals by type and gender from 2003 through 2010 as recorded by DICOSE, calf births as recorded by DICOSE from 2005 – 2010, and

SNIG transaction data from 2006 – 2010. We allow for the possibility that the DICOSE database may be missing data, rather than assuming that all blanks in the data set represent zeroes, but we assume that the SNIG data are complete. While SNIG may not be perfectly accurate with respect to transaction information, non-compliance with SNIG would require both parties to a transaction as well as MGAP inspectors agree to not report a sale and overlook missing paperwork (all animals are checked for proper paperwork before slaughter). Given the level of coordination and potential penalties involved, we think the assumption of completeness of the SNIG data is a reasonable one. Summary statistics for each year of the data set are presented below in Table 3.

Table 3. Pooled Summary Statistics, 2003 – 2010						
	Full Sample		Beneficiaries		Non-beneficiaries	
	Mean	S.D.	Mean	S.D.	Mean	S.D.
DICOSE declaration (Yes=1)	0.90	0.30	0.95	0.21	0.90	0.30
Net calf sales	-0.33	50.12	6.19	31.30	-0.45	50.39
Calf births[a]	30.19	67.85	36.84	45.58	30.07	68.18
Land (ha)	173.23	305.14	152.07	204.47	173.62	306.65
Land (ha), quality weighted	160.42	296.40	121.39	154.14	161.12	298.31
Calves, under 1 year	30.80	66.57	31.16	39.21	30.79	66.96
Heifers, ages 1 to 2	14.54	36.44	15.24	22.13	14.52	36.65
Heifers, ages 2 and up	6.88	22.63	7.03	14.94	6.88	22.74
Breeding cows	51.15	100.50	60.68	69.57	50.98	100.96
Fattened cows (non-breeding)	4.05	21.69	2.64	10.68	4.07	21.83
Steers, ages 1 to 2	12.64	45.30	8.15	20.15	12.72	45.62
Steers, ages 2 to 3	9.23	39.73	5.93	20.17	9.29	39.99
Steers, ages 3 and up	6.50	35.53	3.94	17.78	6.55	35.77
Bulls	1.95	5.83	2.11	3.29	1.95	5.87
Nursing lambs	4.70	42.49	4.34	29.76	4.70	42.69
Male lambs, ages 1 to 2	14.80	61.82	16.24	39.60	14.77	62.15
Female lambs, ages 1 to 2	18.03	58.54	21.87	45.86	17.96	58.75
Lambs, ages 2 to 4	6.47	33.69	6.46	18.56	6.47	33.90
Rams	2.58	10.79	2.91	7.95	2.58	10.83
Wethers	15.18	67.67	14.07	40.75	15.20	68.06
Breeding ewes	70.13	186.80	87.13	145.50	69.82	187.45
Fattened ewes (non-breeding)	4.22	27.69	4.33	26.30	4.22	27.72
Sheep purchases (LU)[b]	1.60	20.48	1.72	9.44	1.60	20.62
Sheep sales (LU)	5.38	33.94	6.42	19.89	5.36	34.14
Cattle sales (LU)	28.54	106.93	23.81	47.12	28.63	107.70
Cattle purchases (LU)	13.01	89.84	9.95	37.59	13.07	90.50
Sheep stock (LU)	27.22	74.62	31.47	54.43	27.15	74.93
Cattle stock (LU)	107.44	202.16	108.26	124.01	107.43	203.30
Observations	23,214		413		22,801	

[a,b] Calf birth data from DICOSE are available beginning in 2005, while SNIG transaction data are available beginning in 2006. Remaining variables run from 2003 through 2010.

Table 3 includes pooled means and standard deviations for the years 2003 – 2010. The first two columns include summary statistics for all cattle producers who reported DICOSE data at least once, with the exception of ULP beneficiaries entering the program in 2006 and 2007, who were filtered out of the database. The next two columns include data for 413 of the 691 producers who had business plans approved in 2008, 388 of which completed their business plans. Beneficiaries that never reported DICOSE data or did not own land in any year from 2003 through 2010 were filtered out of the database by MGAP prior to making it available. Thus our analysis measures ULP impacts on beneficiaries who owned land and reported DICOSE data in at least one year. As indicated in the first row of Table 3, on average 90% of producers listed in the database made a DICOSE declaration each year. However, 32% of all producers and 20% of beneficiaries from the cohorts entering the ULP in 2008 have at least one year of missing data.

The two impact indicators on which we focus our analysis are given in the second and third rows of Table 3, i.e., calf births and net calf sales. As expected, net calf sales in the entire data set are approximately zero; weaned calves are an output for breeders and an input for cattle fatteners, who raise them to be sold to slaughterhouses. In contrast, net calf sales are positive for program beneficiaries, which is what we would expect in a program focused on cattle breeding. These differences are also seen in the number of breeding cows (cows that produce calves) and steers owned by beneficiaries and non-beneficiaries, with the livestock portfolio of beneficiaries tilted more heavily towards the breeding stock. Beneficiaries have less land than non-beneficiaries in absolute and quality-weighted terms, which is not surprising given ULP eligibility requirements. The mean and standard deviation of the land variables indicate the nearly all beneficiaries qualify as small producers, with less than 500 hectares of land. Herd sizes in Livestock Units (LU) are very similar across beneficiaries and non-beneficiaries, while the latter group is more active in terms of cattle transactions. Virtually all of the means in Table 3 are significantly different across beneficiaries and non-beneficiaries, as are there higher order moments and interactions.

5. Estimation method

Selection of ULP beneficiaries (the “treatment group”) was non-random. As a result, we must make assumptions beyond those of a randomized control trial in order to identify the ATT of the ULP on our outcome indicators, calf birth and net calf sales. Let D_i be a dummy variable taking a value of 1 for ULP beneficiaries (the treatment group) and zero for non-beneficiaries (the control group), and suppose the ULP begins in period t with impact realized in period $t + 1$ and beyond. Let CS^1_{it+k} and CS^0_{it+k} represent net calf sales for producer i in period $t + k$ with and without treatment (i.e., with and without participation in the ULP), respectively, where k equals 1 or 2 so that period $t + k$ is either one or two years after the start of treatment. We split pre-program producer characteristics into two vectors: $\mathbf{w}_{it}$, which includes the information collected by DICOSE (the quantity of animals owned, organized by age and gender; quantity and quality of owned land, and calf births beginning in 2005), and $\mathbf{z}_{it}$, which includes the information collected by the SNIG system (sales and purchases of animals, organized by age and gender; this includes net calf sales) and consists of blanks for the years from 2003 – 2005 during which SNIG data are not available. Note that rather than observing $\mathbf{w}_{it}$, we observe $\mathbf{s}_{it}\mathbf{w}_{it}$, where s_{is} is a binary indicator taking a value of 1 if producer i makes a DICOSE declaration at time s, and $\mathbf{s}_{it}$ is a square matrix whose main diagonal contains the history of the DICOSE indicator through period t and is otherwise filled with zeroes.

Our first key identifying assumption is:

(1) $$CS^0_{it+k} \perp D_i \mid \mathbf{s}_{it}, \mathbf{s}_{it}\mathbf{w}_{it}, \mathbf{z}_{it}$$

where $0 \le k \le 1$. This is the Conditional Independence Assumption (CIA), i.e., the independence of the untreated potential outcome from treatment status conditional on observed characteristics (Rosenbaum & Rubin, 1983).

To estimate the ATT without bias, we need an unbiased estimator of the average outcome that would have obtained in the treatment group in the absence of the program, i.e., $E\lfloor CS^0_{it+k} \mid D_i = 1 \rfloor$. We use inverse probability weighting (Wooldridge, 2007). Consider the conditional expectation of observed calf sales for the control group:

(2) $$E\lfloor (1-D_i) CS_{it+k} \mid \mathbf{x}_{it} \rfloor = E\lfloor (1-D_i) CS^0_{it+k} \mid \mathbf{x}_{it} \rfloor = E\left[(1-D_i) \mid \mathbf{x}_{it}\right] E\left[CS^0_{it+k} \mid \mathbf{x}_{it}\right] = \left(1-P\left(D_i = 1 \mid \mathbf{x}_{it}\right)\right) E\left[CS^0_{it+k} \mid \mathbf{x}_{it}\right]$$

where we have collected all observables for producer *i* through time *t* in the *t* x *m* matrix $\mathbf{x}_{it}$. Since $E\left[CS_{it+k}^{0} \mid D_i = 1\right] = E\left\lfloor E\left[CS_{it+k}^{0} \mid \mathbf{x}_{it}\right] P\left(D_i = 1 \mid \mathbf{x}_{it}\right) / P\left(D_i = 1\right)\right\rfloor$, equation (2) implies that we can recover the average of the counterfactual untreated outcome for the treatment group by weighting net calf sales among non-beneficiaries by $P\left(D_i = 1 \mid \mathbf{x}_{it}\right) / \left(\left(1 - P\left(D_i = 1 \mid \mathbf{x}_{it}\right)\right) P\left(D_i = 1\right)\right)$.

To estimate average impacts on calf births, we modify the CIA to account for the presence of missing observations in the DICOSE dataset. We make the following assumption:

(3)
$$S_{it+k}^{1}, S_{it+k}^{0} \perp D_i \mid \mathbf{x}_{it}$$
$$S_{it+k}^{0} CB_{it+k}^{0} \perp D_i \mid \mathbf{x}_{it}$$

where CB_{it+k}^{0} are calf births for producer *i* in year *t* + *k* in the absence of treatment. The implication of the first line of (3) is that the proportion of ULP beneficiaries with characteristics $\mathbf{x}_{it}$ that makes a DICOSE declaration at time *t* + *k* is equal to proportion of non-beneficiaries with the same observed characteristics that would have made a DICOSE declaration if they had been ULP beneficiaries; the same holds with respect to DICOSE declarations without ULP participations. Given that we have data on six years of DICOSE declarations prior to the start of the ULP, this seems like a reasonable assumption. The second line of (3) states that calf births that would be observed in the DICOSE dataset in the absence of treatment, i.e., $S_{it+k}^{0} CB_{it+k}^{0}$, are independent of treatment status conditional on the observed history of each producer. Given what we can observe in the data, this seems a much more reasonable assumption than the alternative that CB_{it+k}^{0}, which we only observe when a DICOSE declaration is made, is independent of treatment status after conditioning on observables.

We make the following additional assumptions regarding the DICOSE declaration indicator, *S*, and calf births, *CB*:

$$(4)\quad \begin{gathered} P\left\lfloor S_{it+k}^{d}=1 \mid H_{it+k}^{d}>0,\mathbf{x}_{it-1}\right\rfloor \geq P\left\lfloor S_{it+k}^{d}=1 \mid H_{it+k}^{d}=0,\mathbf{x}_{it-1}\right\rfloor \\ CB_{it+k}^{d} \perp S_{it+k}^{d} \mid H_{it+k}^{d}>0,\mathbf{x}_{it-1} \\ H_{it+k}^{d}=0 \Rightarrow CB_{i,t+k}^{d}=0 \end{gathered}$$

where H_{it}^{d} is the size of the herd for producer *i* in year *t* given treatment status *d*. The first line states that the conditional probability of making a DICOSE declaration is weakly lower for producers who have zero stock. Given the rules on making livestock declarations, it seems reasonable to assume that producers with positive stock are more likely to make a declaration than producers with zero stock, particularly after conditioning on their observed histories of DICOSE declarations and other information.

The second line of (4) states that within groups of farmers with the same observed characteristics through period *t* – 1, the same treatment status, and positive herd sizes at time *t*, there is no dependence between the potential outcomes for calf births and the decision to make a livestock declaration. We had previously assumed that producers who have followed the same dynamics for calf births up to a random shock through time *t* would continue to do so through period *t* + 1. Now we are also assuming that any random shocks to calf births cannot affect the decision to make a DICOSE declaration, conditional on having a positive herd size. The third line of (4) states that producers with a zero herd size at the time of DICOSE declarations must not have had any calf births since the time of the last DICOSE declaration. Calves are usually weaned after six months, making it unlikely that all calves born between annual DICOSE declarations would have been sold, and a calf mortality rate of 100% is not realistic.

In the appendix, we show that the assumptions given in (3) and (4) are sufficient to partially identify the ATT of the ULP on calf births. The bounds on the ATT conditional on observed characteristics $\mathbf{x}_{it-1}$ are:

$$E\left[S_{it}^{1}CB_{it}^{1} \mid D_i = 1, \mathbf{x}_{it-1}\right] - E\left[\frac{\left(1-D_i\right)S_{it}^{0}CB_{it}^{0}P\left(D_i = 1 \mid \mathbf{x}_{it-1}\right)}{\left(1-P\left(D_i = 1 \mid \mathbf{x}_{it-1}\right)\right)P\left(D_i = 1\right)P\left(S_{it}^{0} = 1 \mid \mathbf{x}_{it-1}\right)} \mid \mathbf{x}_{it-1}\right]$$

(5)
$$\leq E\left[CB_{it}^{1} \mid \mathbf{x}_{it-1}, D_i = 1\right] - E\left[CB_{it}^{0} \mid \mathbf{x}_{it-1}, D_i = 1\right] \leq$$

$$\frac{E\left[S_{it}^{1}CB_{it}^{1} \mid D_i = 1, \mathbf{x}_{it-1}\right]}{P\left(S_{it}^{1} = 1 \mid D_i = 1, \mathbf{x}_{it-1}\right)} - E\left[\frac{\left(1-D_i\right)S_{it}^{0}CB_{it}^{0}P\left(D_i = 1 \mid \mathbf{x}_{it-1}\right)}{\left(1-P\left(D_i = 1 \mid \mathbf{x}_{it-1}\right)\right)P\left(D_i = 1\right)} \mid \mathbf{x}_{it-1}\right]$$

The bounds can be estimated using weighted averages of the observed outcomes among the treated and untreated producers.

The ATT on calf sales in 2009 and 2010 is estimated as:

(6)
$$ATT_{CS_{it+k}} = \frac{\sum_{i=1}^{N} D_i CS_{it+k}}{\sum_{i=1}^{N} D_i} - \frac{\sum_{i=1}^{N} \omega_i^{CS}\left(1-D_i\right)CS_{it+k}}{\sum_{i=1}^{N} \omega_i^{CS}\left(1-D_i\right)}$$

where $t + k$ now equals either 2009 or 2010. We normalize by the sum of the weights to avoid arbitrarily large weights (Busso, DiNardo, & McCrary, 2011). The expression for the weights is:

(7)
$$\omega_i^{CS} = D_i + \left(1-D_i\right)\frac{\hat{p}\left(\mathbf{x}\right)}{\left(1-\hat{p}\left(\mathbf{x}\right)\right)\rho}$$

$$\hat{p}\left(\mathbf{x}\right) = \frac{\exp\left(\mathbf{x}_{i2008}'\hat{\boldsymbol{\gamma}}\right)}{1+\exp\left(\mathbf{x}_{i2008}'\hat{\boldsymbol{\gamma}}\right)}, \rho = \frac{\sum_{i=1}^{N} D_i}{N}$$

where $\hat{p}\left(\mathbf{x}\right)$ is the propensity score for selection into ULP, ρ is our estimate of $P\left(D_i = 1\right)$ and $\mathbf{x}_{i2008}$ contains producer i's observed characteristics from 2003 through 2008. The elements of $\mathbf{x}_{i2008}$ include calves, breeding cows, total stock of cows, total stock of steers, and total stock of sheep (all in livestock units) declared to DICOSE each year from 2003 through 2008, the first component of a principal components decomposition of quality-weighted land from 2003 through 2008,[3] an indicator for making a DICOSE declaration in each year for 2003 through 2007,[4] calf births for 2005 through 2008, and net sales of calves, other cows, and sheep in livestock units for 2006 through 2008. We also included the squares of all continuous variables and the interaction between all variables in each year, as well as the interactions across years for calf sales and calf births both with themselves and with each other. This logit specification should allow us to balance the observed pre-ULP means, variances, and covariances of all variables across the treatment and control groups, as well as the variance and covariance of calf births and calf sales across years.

To estimate the bounds for the ATT on calf births in 2009 and 2010, we define the following weights:

(8)
$$\begin{aligned}\omega_{it+k}^{lb} &= D_i + \left(1-D_i\right)\frac{\hat{p}\left(\mathbf{x}\right)}{\left(1-\hat{p}\left(\mathbf{x}\right)\right)\rho\hat{p}\left(S_{it+k}^{0}\right)}\\ \omega_{it+k}^{ub} &= \frac{D_i}{\hat{p}\left(S_{it+k}^{1}\right)} + \left(1-D_i\right)\frac{\hat{p}\left(\mathbf{x}\right)}{\left(1-\hat{p}\left(\mathbf{x}\right)\right)\rho}\end{aligned}$$

where $t + k$ is 2009 or 2010, $\hat{p}\left(S_{it+k}^{0}\right)$ is the estimated probability of making a livestock declaration at time t without treatment (estimated using the control group), and $\hat{p}\left(S_{it+k}^{1}\right)$ is the estimated probability of making a livestock at time $t + k$ declaration with treatment (estimated using the treatment group). The fitted values $\hat{p}\left(S_{it+k}^{0}\right)$ are generated from another logit regression using the same variables as the model for $\hat{p}\left(\mathbf{x}\right)$ as described above, but are estimated using just the controls. The fitted values $\hat{p}\left(S_{it+k}^{1}\right)$ are estimated only for 2010; all but two beneficiaries made a DICOSE declaration in 2009, and these two are dropped from the dataset. The model for $\hat{p}\left(S_{it+k}^{1}\right)$ uses a subset of $\mathbf{x}_{i2008}$, including calf births, calves, breeding cows, steers, and sales of calves, other cattle, and

[3] The first component explained 92% of the variation in quality-weighted land from 2003 through 2008.
[4] We do not include an indicator for making a DICOSE declaration in 2008 because all ULP beneficiaries made a declaration that year.

sheep for 2007 and 2008. The smaller subset of explanatory variables was chosen after examining the balance in observed characteristics between the 13 treatment group members who did not make a DICOSE declaration in 2010 and the 400 who did, and taking into consideration the smaller size of the treatment group relative to the controls.

The bounds on the ATT for calf births in period $t + k$ are estimated as:

$$(9) \qquad \left[ATT_{CB_{it+k}}^{lb}, ATT_{CB_{it+k}}^{ub} \right] = \left[\begin{array}{c} \dfrac{\sum_{i=1}^{N} \omega_{it+k}^{lb} D_i S_{it+k} CB_{it+k}}{\sum_{i=1}^{N} \omega_{it+k}^{lb} D_i} - \dfrac{\sum_{i=1}^{N} \omega_{it+k}^{lb} \left(1 - D_i\right) S_{it+k} CB_{it+k}}{\sum_{i=1}^{N} \omega_{it+k}^{lb} \left(1 - D_i\right)}, \\ \dfrac{\sum_{i=1}^{N} \omega_{it+k}^{ub} D_i S_{it+k} CB_{it+k}}{\sum_{i=1}^{N} \omega_{it+k}^{ub} D_i} - \dfrac{\sum_{i=1}^{N} \omega_{it+k}^{ub} \left(1 - D_i\right) S_{it+k} CB_{it+k}}{\sum_{i=1}^{N} \omega_{it+k}^{ub} \left(1 - D_i\right)} \end{array} \right]$$

The lower bound on the ATT over both post-ULP years is estimated by averaging the lower bounds for 2009 and 2010, and the upper bound is estimated similarly.

6. Inference

Statistical inference for the ATT parameters of interest is complicated by two factors: estimated variances of program impacts should be corrected for their dependence on the estimated parameters of the logit regressions, and the ATT on calf births is not point-identified. Note that the ATT for calf sales for 2009 and 2010 as well as the bounds on the ATT for calf births in 2009 and 2010 are all weighted differences in means comparing ULP beneficiaries and non-beneficiaries, and are therefore identical to the coefficient on D_i in a weighted regression of calf sales or calf births on a column of ones and D_i. We can therefore estimate the propensity score parameters (for selection into the ULP and into the dataset in 2009 or 2010) as well as the ATT parameters in a system of equations consisting of the moment conditions for the various logit models and weighted least squared regressions.

The moment conditions give rise to a straightforward "sandwich" estimator of the covariance matrix, i.e., a matrix of the form $\mathbf{A}^{-1}\mathbf{BB}'\mathbf{A}^{-1}/N$, where $\mathbf{BB}'$ is the outer product of the moment conditions and $\mathbf{A}^{-1}$ is the inverse of the matrix of second order conditions of each moment condition with respect to all other model parameters. The resulting covariance matrix is consistent and robust to heteroscedasticity under standard regularity conditions, and it incorporates the uncertainty in the weights into the standard errors for the ATT parameters (Wooldridge, 2002). We scale the estimated covariance matrix by a degrees of freedom correction $N/(N-G)$, where G is the total number of model parameters.

To account for the fact that the ATT on calf births in 2009 and 2010 is partially identified, we use the technique of Imbens and Manski (2004). They derive $\sqrt{N}$ -consistent, 95% confidence intervals for partially identified parameters when the bounds on the parameter are functions of averages of the observed data and probabilities of selection into the observed data. In the present context, their formula is given by:

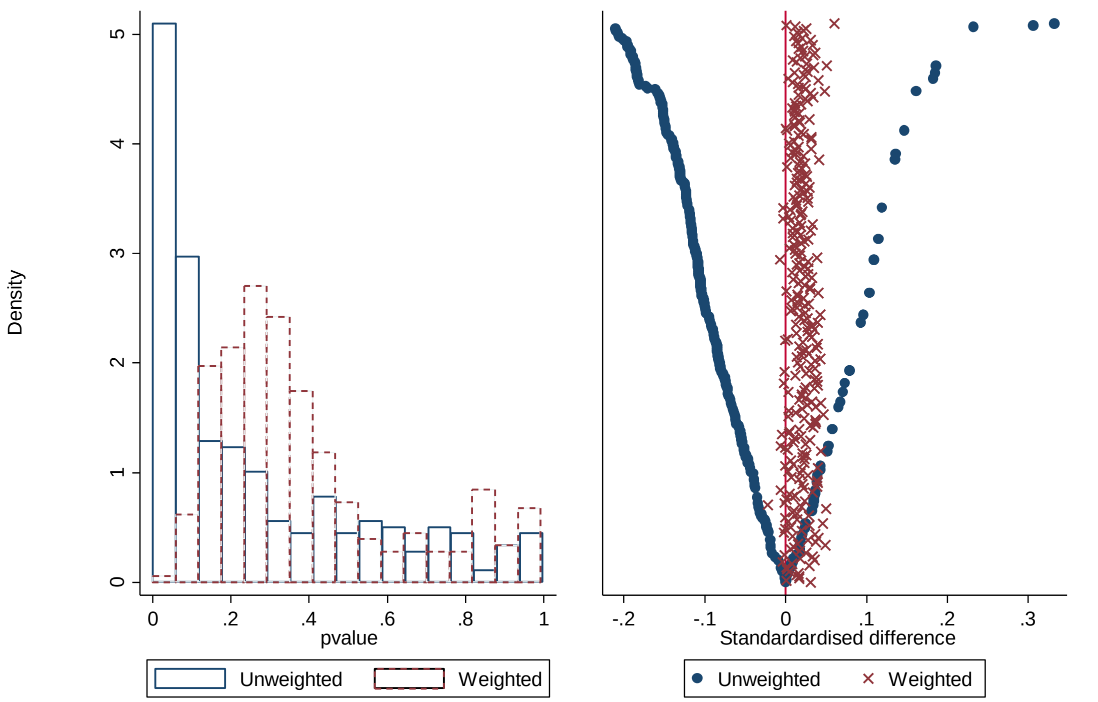
(10)
$$\left\lfloor ATT^{lb}_{CB_{it+k}} - \bar{C}_N \hat{\sigma}_{ATT^{lb}_{t+k}}, ATT^{ub}_{CB_{it+k}} - \bar{C}_N \hat{\sigma}_{ATT^{ub}_{t+k}} \right\rfloor, \text{ where}$$
$$\Phi\left(\bar{C}_N + \frac{ATT^{ub}_{CB_{it+k}} - ATT^{lb}_{CB_{it+k}}}{\max\left(\hat{\sigma}_{ATT^{lb}_{t+k}}, \hat{\sigma}_{ATT^{ub}_{t+k}}\right)}\right) - \Phi\left(-\bar{C}_N\right) = 0.95$$

Φ is the cumulative standard normal distribution, and $\hat{\sigma}_{ATT^{lb}_{t+k}}$ and $\hat{\sigma}_{ATT^{ub}_{t+k}}$ are the estimated standard errors for the lower and upper bounds of the ATT on calf births in period $t + k$, and are taken from our estimate of the model parameter covariance matrix.

7. Covariate balance with and without inverse propensity score weighting

The assumptions used here to identify and bound the ATT of the ULP require that we can create a comparison group that is observationally identical to the treatment group prior to the start of the ULP. Figure 1 depicts the improvement in covariate balance that results from weighting covariates using the weights given in equation (7).

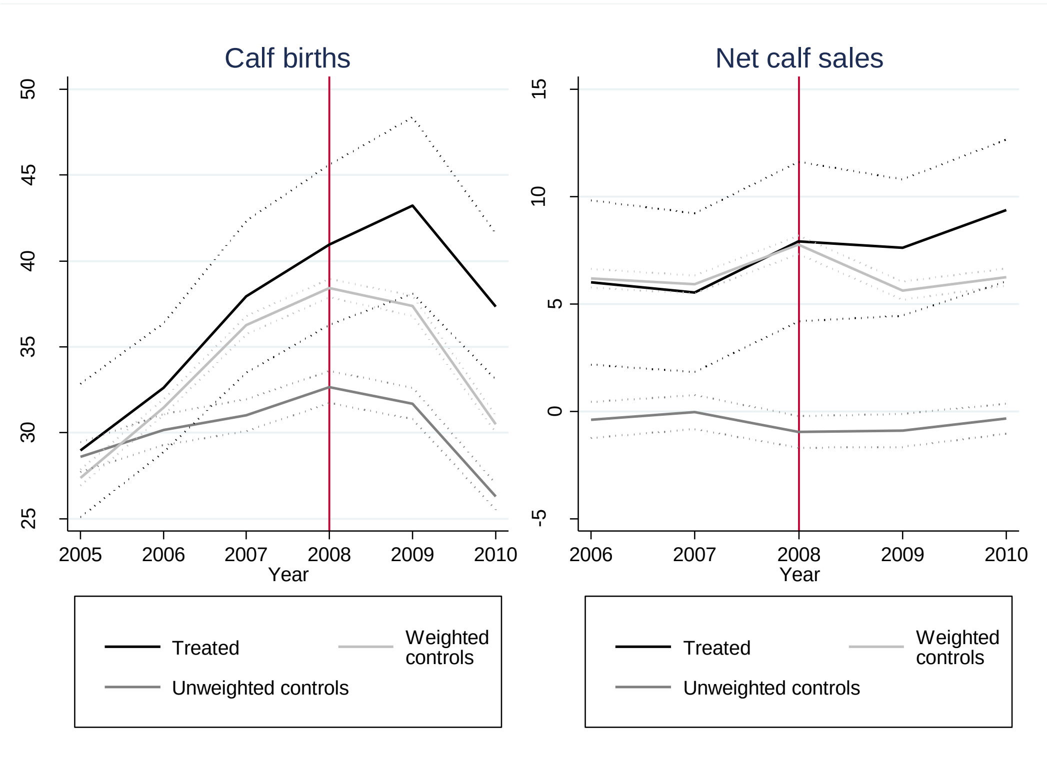
Figure 1. Covariate balance



The left panel of Figure 1 depicts the distribution of p-values from t-tests comparing the means in the treatment and control groups of all the variables used in estimation of $\hat{p}\left(D_i\right)$, $\hat{p}\left(S^1_{it+k}\right)$, and $\hat{p}\left(S^0_{it+k}\right)$, including all interactions. Prior to weighting, the bulk of the mass of the distribution of p-values is in the left tail, while weighting shifts the mass to the right. Initially, the vast majority of differences are significant, whereas none are after weighting. Since covariate balance is an in-sample property rather than a population property, t-tests are a flawed tool with respect to judging covariate balance, as they are sensitive to sample size. Therefore we include the right panel of Figure 1, which shows the standardized differences (the differences in means across groups as a proportion of the standard deviation of the variable in the sample) of all model covariates with and without weighting; the right panel was created using the user-written Stata program *pbalchk*. The right panel shows that balance is sharply improved for nearly all covariates. Covariates for which we observe no change or a worsening of balance are those for which there was virtually no difference across the treatment and control groups prior to weighting.

The validity of our assumptions also hinge whether we can create a comparison group whose outcome dynamics are identical to those of the treatment group prior to the start of the ULP in 2008. Figure 2 shows the dynamics of calf sales and observed calf births both with and without weighting.

Figure 2. Observed dynamics of calf births and calf sales



In the left panel, the topmost dark line shows average calf births as reported in DICOSE for the treatment group, both before and after the start of the ULP, and the dotted lines bracketing it are its 95% confidence interval (computed using the standard normal distribution). The lowest gray line depicts average calf births for the controls prior to weighting, while the line in the middle shows controls after weighting. All groups are following an upward trend prior to the start of the program, likely reflecting recovery of the livestock sector from an outbreak of hoof and mouth disease in 2001, and calf production drops in 2009 and 2010, reflecting the effects of a severe drought that impacted pasture availability.

Despite these general observations, the treatment and control groups are clearly following different trends prior to the start of the ULP without weighting, and the differences are significant in 2007 and 2008. Once we use weights, the trends follow each other closely, and do not become significantly different until after the start of the ULP. The right panel tells a similar story for net calf sales, where we see significantly different paths for the treatment and control groups without weighting, identical paths through 2008 once we use the propensity score weights, and divergence starting in 2009.

8. Results

Table 4 below reports the estimated ATTs for calf sales in 2009 and 2010, the overall ATT on calf sales across both years, and the bounds on the ATT for calf births in 2009 and 2010 as well as bounds on the overall effect:

Table 4. ATT for calf sales, 2009-2010

Year	ATT	SE	t-statistic	p-value	95% Confidence Interval
Total	4.35	2.26	1.93	0.054	[-0.07, 8.77]
2010	2.64	1.51	1.75	0.08	[-0.31, 5.59]
2009	1.71	1.45	1.18	0.24	[-1.14, 4.56]

The first row gives the overall ATT for calf sales, i.e., combining impacts over 2009 and 2010. From 2009 through 2010, ULP beneficiaries sold an average of 4.35 calves annually more than they would have had they not participated in the program. When we disaggregate the impact by year, we see that the 2010 impact is estimated more precisely than the impact in 2009, while both are less precise than the overall impact. The greater precision of the overall effect is due to the negative estimated covariance between the ATTs for each year.

Table 5 displays the estimated bounds and 95% confidence intervals for the overall ATT on calf births as well as the ATT on calf births in 2009 and 2010.

Table 5. Bounds for ATT on calf births, 2009-2010

	Total	2010	2009
Bounds	(11.36, 15.30)	(5.35, 7.91)	(6.02, 7.39)
95% CI	[4.53, 22.39]	[1.28, 11.16]	[1.45, 13.02]

The results in Table 5 indicate that from 2009 through 2010, ULP beneficiaries produced between 11.36 and 15.3 more calves on average than they would have had they not participated in the program. This overall effect is not driven by the impact in any particular year, as the bounds for the ATT in 2009 and 2010 are quite similar and the ATT in each year is significant at a 5% level.

Translating these estimated impacts into a net social benefit requires calculating the IRR of the program. Our data are missing several key pieces of information needed to construct a thorough measure of IRR, such as the portion of total program costs attributed to producer business plans and producer profits. In addition, any measure of IRR constructed using our data will be limited to beneficiaries who owned land. Instead, we calculate what we believe to be a reasonable approximation to IRR given the available data. We assume that on average the subsidy covered half the cost of each business plan. We also assume that the all business plan costs represented increases relative to the no program scenario, and thus take our measure of cost to be twice the average subsidy received by program beneficiaries. We calculate the gross value of program impacts using the estimation results reported above and national-level livestock price data for the internal Uruguayan market as provided by the National Meat Institute (INAC). Since the year 2000, the data showed no significant upward trend in the real price per kilogram for calves (we focus on the post 2000 period because the 1990s were a period of major economic reforms); thus we use the average price per kilogram of calf meat from 2000 – 2011 of $1.13 in calculating the gross value of ULP impacts. Since our data do not include animal weights, we assume a weight of 140 kilograms per animal; this is the weight recommended by MGAP for permanent weaning.

Table 6. Internal rate of return

Lifespan of benefits (beyond 2010)	10	15	20	25

IRR value of production	(6%, 19%)	(12%, 25%)	(15%, 27%)	(17%, 28%)
IRR net calf sales	-13%	-4%	0%	3%

The results in Table 6 suggest that the IRR is sensitive to how benefits are measured. If benefits are measured in terms of the value of additional production, as given by the bounds presented in the first row of Table 6, they are quite clearly positive and large in magnitude. Note that these benefits do not include the cost of maintaining the additional calves produced, e.g., the opportunity cost of pasture and labor costs. However, when using net calf sales benefits are far more modest. The useful lifespan of knowledge gained and investments made as a result of the ULP would have to be at least 25 years in order for the IRR to be equal to 12%, a figure that could be considered a reasonable return on the resources put into the program. On the other hand, the severe drought that affected Uruguay lowered the demand for calves, as feeders did not have adequate pasture to support their usual level of demand. This may have blunted the impacts of the ULP on calf sales, suggesting that the IRR calculated above using net calf sales is very conservative. We conclude somewhat cautiously that the evidence in Table 6 suggest the ULP was a modest success in terms of net benefits.

9. Robustness checks

Prior to discussing our results in detail, we conduct robustness checks of our estimation method by estimating "placebo" treatment effect, i.e., we estimate the ATT for pre-treatment years. First, we use data through 2006 to estimate the weights used to compute the ATTs, and estimate placebo ATT parameters for 2007 and 2008. We then check to see if we can reject the null hypothesis that the ATT is zero, knowing that the ATT should be zero since the ULP was not yet operating. For the ATT on calf sales, we also check the bias reduction of our probability-weighted estimator versus a naïve comparison of ULP beneficiaries and non-beneficiaries in 2007 and 2008, i.e., the ATT estimator we would use had ULP participation been randomly assigned. We then repeat this exercise just for 2008, using data through 2007 to estimate the probability weights, in order to see how the performance of the estimator changes as more information is added to the conditioning set and the model only needs to extrapolate out the dynamics of the outcomes for a single period rather than two. Table 7 below shows the results of our robustness checks for the estimated ATTs of calf sales.

Table 7. Placebo ATTs, calf sales

Year	Conditioning set	Naïve estimate	Estimated ATT	Bias reduction	SE	T-stat	p-value
2008	2003 – 2006	6.41	2.81	56%	1.22	2.31	0.02
2007	2003 – 2006	4.02	0.37	92%	1.39	0.22	0.82
2008	2003 – 2007	6.41	1.31	79%	1.26	1.04	0.30

It is important to remember that sales data are only available beginning in 2006. As a result, the first two rows of Table 7 show estimated placebo ATTs on calf sales in 2007 and 2008 after controlling for a single year of calf sales data, as well as the history of other predictors since 2003. This is demanding much more of the data than would be the case when estimating the ATTs for 2009 and 2010, where we are able to condition on sales and purchase data for 2006 through 2008. The third row gives the placebo ATT for calf sales in 2008 after controlling for observables through 2007. It seems clear that the model has much more success eliminating bias at time t when we hold observables fixed through time $t - 1$ and less success eliminating bias at time $t + 1$. This should serve as a caveat when interpreting the results we present in Table 4. Also note that the precision of the estimated placebo ATTs in Table 7 is similar to that of the actual estimated ATTs in Table 4; failure to reject the null hypothesis of no effect when estimating placebo ATTs is not the result of lower precision caused by using few years of data.

Next we estimate bounds for placebo ATTs on calf births in 2007 and 2008, as shown in Table 8.

Table 8. Bounds on placebo ATTs, calf births

Conditioning set	2003 - 2006		2003 – 2007
Year	2008	2007	2008
Bounds	(5.73, 6.08)	(-31.37, 11.92)	(3.09, 3.89)
95% CI	[1.71, 10.11]	[-35.09, 15.98]	[-0.53, 7.51]

If the model is successful in removing bias in the estimated bounds of the ATT on calf births, the confidence intervals for the true parameter in each placebo test should contain zero. As in the case of calf sales, we see that the model is more successful in removing bias one year after the end of the time horizon contained in the conditioning set; when controlling for observables through 2006, we cannot reject the null hypothesis of no effect in 2007 (and the bounds themselves contain zero), and when controlling for observables through 2007 we cannot reject the null hypothesis of no effect in 2008. Note that just as in the case of the estimated placebo ATTs for calf sales, failure to reject the null in Table 8 is not being driven by a loss in precision due to using fewer years of data. We conclude that our estimated ATTs and ATT bounds for 2009 are more credible than those for 2010.

10. Conclusion

This paper evaluated the impact of the ULP, a publicly funded, privately delivered extension program for cattle breeders in Uruguay. Focusing on cattle breeders that reported owning some land from 2003 to 2008, we found significant impacts on both calf production and sales of calves minus purchases in 2009 and 2010. Robustness checks suggest that our 2009 results are more credible. Therefore the most conservative conclusion one can draw from our paper is that the program had significant impacts on calf production in 2009.

Our results suggest that the program had the desired effect in terms of sign but that the magnitudes of its impacts were modest, leading one to ask what characteristics of the ULP and cattle breeding in Uruguay were driving the size of program effects. The design of the ULP is based on the belief that a lack of knowledge about sound management techniques was leading to suboptimal outcomes among cattle breeders, and that private extension agents could be an effective means of delivering this information. For this to be the case, it must have been that producers faced constraints which prevented them from obtaining information on good management practices on their own or from extension agents outside the context of the ULP. The public extension service in Uruguay is small, with around 500 agents, and while 52% of producers have used extension at some point, the usage rate is highly correlated with land holdings and dominated by use of veterinary services (MGAP 2002). In addition, the flow of information on management practices is likely hampered in the case of Uruguay by the geographic isolation of livestock producers, who tend to interact only with a small number of producers in their immediate vicinity for the purpose of labor exchanges (Morales, et al., 2010). Higher use of mobile phones and the internet may weaken the effects of isolation, but geography appears to be a relevant constraint for the time period considered in this evaluation.

Constraints on the flow of information therefore seem plausible. However, available data makes it impossible to test this assumption with any rigor, and MGAP would have faced the same scarcity of relevant data prior to the start of the ULP. The information collected through the agricultural census and the NLS indicate that management practices were poor and use of extension was lower among small and medium sized producers, but neither instrument collected the information needed to identify binding constraints. The evidence that is available is based on interviews with small numbers of cattle breeders, making it difficult to reach conclusions about the population of cattle breeders overall.

In addition to addressing constraints caused by geographic isolation and the limited reach of public extension, the subsidy offered to program participants and used to compensate extension agents may have loosened liquidity on producers. On the supply side, the subsidy may have made it more profitable for extension agents to work with small and medium producers; in the absence of a subsidy, the transaction costs associated with working with a large number of small producers can lead to a low supply of private extension to everyone but the largest producers (Anderson & Feder, 2004). It seems doubtful that the subsidy could have had such an effect, however. Firstly, the subsidy was paid in dollars, and the exchange rate of the peso to the dollar climbed from 0.039 in January 2007 to a peak of 0.051 in August 2008, fluctuating between 0.041 and 0.051 for the 30 months over which subsidies were paid out to the cohort entering the ULP in

2008.[5] The appreciation of the peso meant that a beneficiary entering the program in July 2008 would receive 29% less from the first payment than would have been the case if he had entered the ULP one year earlier. In addition, private extension workers typically receive around $150 per day for working with a producer, suggesting that a total payment from a beneficiary to an extension agent of $320 would not provide a strong incentive on its own. There are anecdotal accounts of producers and extension agents agreeing to a more equitable split of the subsidy, but the available data do not allow us to quantify this phenomenon (Bentancur, Fernández, Rado, & Zurbriggen, 2012).

The same anecdotal evidence describing additional compensation to extension agents suggests that the main motivation for extension agents was using the program as a means to find new clients. Furthermore, uncertainty over the value of extension advice is sometimes cited as a reason why demand for private extension services is low among small and medium producers (Feder, Willett, & Zijp, 1999). The fact that extension agents were approved by MGAP before working with ULP beneficiaries may have reduced uncertainty on the part of producers regarding the value of extension, although evidence suggests that the training given to extension agents as part of the certification process was weak (Rearte, 2011). In any case, incentives to experiment with extension would have been increased by lowering information costs associated with locating an agent and through the subsidy, while the program lowered transaction costs for extension agents by providing a convenient pool of potential clients.

These observations underscore the importance of conducting a rigorous diagnostic prior to policy intervention. In the case of management practices, what appears to be suboptimal behavior on the surface may be an optimal response given market conditions and constraints. By identifying the constraints that bind on producer behavior, and examining how the incidence of different constraints varies with producer characteristics, policy design can be better informed and potentially more successful.

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[5] Exchange rates taken from http://www.oanda.com/currency/historical-rates/.

Appendix: derivation of bounds on the ATT for calf births

In what follows, we derive the bounds on the ATT for calf births given in (5). While we cannot point-identify the ATT of the ULP on calf births, we can use inverse probability weighting to identify bounds on the ATT. Consider the conditional expectation of the observed outcome for the treated group in a post-treatment period *t*:

(A.1)

$$E\left\lfloor S_{it}CB_{it} \mid D_i = 1, \mathbf{x}_{it-1} \right\rfloor =$$
$$E\left[S^1_{i,t}CB^1_{i,t} \mid D_i = 1, \mathbf{x}_{it-1} \right] =$$
$$E\left[S^1_{it}CB^1_{it} \mid \mathbf{x}_{it-1}, H^1_{it} > 0 \right] P\left(H^1_{it} > 0 \mid D_i = 1, \mathbf{x}_{it-1} \right) =$$
$$P\left(S^1_{it} = 1 \mid D_i = 1, \mathbf{x}_{it-1}, H^1_{it} > 0 \right) E\left[CB^1_{it} \mid D_i = 1, \mathbf{x}_{it-1}, H^1_{it} > 0 \right] P\left(H^1_{it} > 0 \mid D_i = 1, \mathbf{x}_{it-1} \right) =$$
$$P\left(S^1_{it} = 1 \mid D_i = 1, \mathbf{x}_{it-1}, H^1_{it} > 0 \right) E\left[CB^1_{it} \mid D_i = 1, \mathbf{x}_{it-1} \right]$$

The second equality follows from the third line of (4), i.e., $H^1_{it} = 0 \Rightarrow CB^1_{it} = 0$.

The third equality follows from the second line of (4).

Equation (A.1) implies that if we could weight $S_{it}CB_{it}$ in the treated group by:

(A.2)
$$\frac{1}{P\left(S^1_{it} = 1 \mid D_i = 1, \mathbf{x}_{it-1}, H^1_{it} > 0 \right)}$$

we could recover the expected outcome conditional on treatment for the treated group.

The term $P\left(S^1_{it} = 1 \mid D_i = 1, \mathbf{x}_{it-1}, H^1_{it} > 0 \right)$ is not point identified by the assumptions above, but it can be bounded:

(A.3)
$$P\left(S^1_{it} = 1 \mid D_i = 1, \mathbf{x}_{it-1} \right) \leq P\left(S^1_{it} = 1 \mid D_i = 1, \mathbf{x}_{it-1}, H^1_{it} > 0 \right) \leq 1$$

The lower bound follows from our assumption that producers with zero stock have a weakly lower conditional probability of making a DICOSE declaration as compared to producers with positive stock. As a result, the conditional probability of making a declaration at time *t* for the whole population of producers with characteristics $\mathbf{x}_{it-1}$ is dragged down by those who have zero stock. The upper bound is obtained when there is perfect compliance with livestock declarations and all missing values are the result of producers not having any livestock. These bounds imply:

(A.4) $$E\left[S_{it}^{1}CB_{t}^{1} \mid D_{i}=1,\mathbf{x}_{it-1}\right] \leq E\left[CB_{it}^{1} \mid D_{i}=1,\mathbf{x}_{it-1}\right] \leq \frac{E\left\lfloor S_{it}^{1}CB_{it}^{1} \mid D_{i}=1,\mathbf{x}_{it-1}\right\rfloor}{P\left(S_{it}^{1}=1 \mid D_{i}=1,\mathbf{x}_{it-1}\right)}$$

Note that $S_{it}^{1}=S_{it}$ and $CB_{it}^{1}=CB_{it}$ for the treatment group so that the upper and lower bounds can be estimated using the data.

Now we turn to the counterfactual untreated average outcome for the treatment group in post-treatment period *t*. Consider the conditional expectation of the observed outcome among the untreated producers in the population:

(A.5) $$\begin{gathered} E\left\lfloor (1-D_{i})S_{it}CB_{it} \mid \mathbf{x}_{it-1}\right\rfloor = \\ \left(1-P\left(D_{i}=1 \mid \mathbf{x}_{it-1}\right)\right)E\left[S_{it}^{0}CB_{it}^{0} \mid \mathbf{x}_{it-1}\right] = \\ \left(1-P\left(D_{i}=1 \mid \mathbf{x}_{it-1}\right)\right)E\left[S_{it}^{0}CB_{it}^{0} \mid \mathbf{x}_{it-1},H_{it}^{0}>0\right]P\left(H_{it}^{0}>0 \mid \mathbf{x}_{it-1}\right) = \\ \left(1-P\left(D_{i}=1 \mid \mathbf{x}_{it-1}\right)\right)P\left(S_{it}^{0}=1 \mid \mathbf{x}_{it-1},H_{it}^{0}>0\right)E\left[CB_{it}^{0} \mid \mathbf{x}_{it-1},H_{it}^{0}>0\right]P\left(H_{it}^{0}>0 \mid \mathbf{x}_{it-1}\right) = \\ \left(1-P\left(D_{i}=1 \mid \mathbf{x}_{it-1}\right)\right)P\left(S_{it}^{0}=1 \mid \mathbf{x}_{it-1},H_{it}^{0}>0\right)E\left[CB_{it}^{0} \mid \mathbf{x}_{it-1}\right] \end{gathered}$$

The second line follows from the second line of the assumptions given in (3), i.e., that the untreated calf births that would be observed in the DICOSE data set are independent of treatment status conditional on observables. The third line expands $E\left\lfloor S_{it}^{0}CB_{it}^{0} \mid \mathbf{x}_{it-1}\right\rfloor$ into its expectations conditional on the sign of H_{it}^{0}, using the assumption that $H_{it}^{1}=0 \Rightarrow CB_{it}^{1}=0$. The fourth line follows from the second line of (4), i.e., making a livestock declaration is independent of the calf births conditional on having a positive herd size and the history of observed characteristics through time $t-1$.

Note that:

(A.6) $$\frac{E\left[CB_{it}^{0} \mid \mathbf{x}_{it-1}\right]P\left(D_{i}=1 \mid \mathbf{x}_{it-1}\right)}{P\left(D_{i}=1\right)} = E\left[CB_{it}^{0} \mid D_{i}=1,\mathbf{x}_{it-1}\right]$$

Together, equation (A.5) and (A.6) imply that we could recover an unbiased estimate of the untreated outcome for the treatment group by weighting $(1-D_i)S_{it}CB_{it}$ by:

(A.7) $$\frac{P(D_i=1|\mathbf{x}_{it-1})}{(1-P(D_i=1|\mathbf{x}_{it-1}))P(D_i=1)P(S_{it}^0=1|\mathbf{x}_{it-1},H_{it}^0>0)}$$

The term $P(S_{it}^0=1|\mathbf{x}_{it-1},H_{it}^0>0)$ in the denominator is not point identified but can be bounded. We have that:

(A.8) $$P(S_{it}^0=1|\mathbf{x}_{it-1})\le P(S_{it}^0=1|\mathbf{x}_{it-1},H_{it}^0>0)\le 1$$

The logic behind (A.8) is identical to that of (A.3). Since we have assumed that making a DICOSE declaration is conditionally independent of treatment status, the lower bound is equal to $P(S_{it}=1|\mathbf{x}_{it-1},D_i=0)$ and can be estimated using untreated producers.

Taken together we have the following bounds on $E\lfloor CB_{it}^0|\mathbf{x}_{it-1},D_i=1\rfloor$:

(A.9) $$E\left[\frac{(1-D_i)S_{it}^0CB_{it}^0P(D_i=1|\mathbf{x}_{it-1})}{(1-P(D_i=1|\mathbf{x}_{it-1}))P(D_i=1)}|\mathbf{x}_{it-1}\right]$$
$$\le E\left[CB_{it}^0|\mathbf{x}_{it-1},D_i=1\right]\le$$
$$E\left[\frac{(1-D_i)S_{it}^0CB_{it}^0P(D_i=1|\mathbf{x}_{it-1})}{(1-P(D_i=1|\mathbf{x}_{it-1}))P(D_i=1)P(S_{i,t}^0=1|\mathbf{x}_{it-1})}|\mathbf{x}_{it-1}\right]$$

Combining the results in (A.4) and (A.9) yields the formulation given in (5).

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