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Free prescriptions for low-income pensioners? The cost of returning to free-of-charge drugs in the Spanish National Health Service

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Abstract

This study estimated the impact of reducing a capped low coinsurance rate for outpatient medicines to nil for low-income pensioners and disabled individuals in the Valencian Community (Spain). This reduction was implemented in January 2016 as a regional reform which modified the national cost-sharing reform adopted in July 2012. The impact of this intervention on the number of monthly prescriptions dispensed between July 2012 and December 2018 was estimated using two different approaches of the synthetic control method, the classical method and the method based on Bayesian structural time series. The estimates from both methods were similar, showing significant overall increases of 6.34% and 6.70% [95% credible interval: 4.05, 9.47], respectively in the number of prescriptions dispensed in this region. These results are similar to those of the previous studies indicating that reducing price from a small amount to zero discontinuously boosts demand. This evidence indicates that the impact of this intervention on the budget of the regional health service is far greater than the amount of the subsidy in the public budget. These results are useful for making accurate budgetary projections for similar eliminations of charges for low-income pensioners in the Spanish National Health Service.

KEYWORDS
cost-sharing, healthcare expenditure, pharmaceuticals, synthetic control

JEL CLASSIFICATION
G22, I18, J14

1 | INTRODUCTION

In June 2012, Spain enacted a reform of the pharmaceutical coinsurance scheme, in which cost-free dispensing of all pensioners' medicines was replaced by a 10% co-payment subject to an €8 monthly cap (for incomes of less than €18,000 per year; Royal Decree Law 16/2012). After the national reforms of 2012, there was a temporary decrease in the number of prescriptions for outpatients, estimated at more than 10% (Antoñanzas, Rodríguez-Ibeas, Juarez-Castelló, & Lorente, 2014; Hernández-Izquierdo, López-Valcárcel, Morris, Melnychuk, & Abásolo, 2019; Puig-Junoy, Rodríguez-Feijoó, &

Lopez-Valcarcel, 2014). However, there were signs that the short-term effect of the intervention was not permanent (Puig-Junoy, Rodríguez-Feijóo, GonzálezLópez-Valcárcel, & Gómez-Navarro, 2016).

In January 2016, the Valencian Community (VC) implemented a regional subsidy[1] which allowed the abolition of charges for low-income pensioners (those with incomes of less than €18,000 per year) and disabled individuals. This method of returning to the system of free prescriptions in the VC was extended to cover children (in February 2017), nonexempted unemployed people (in May 2018), and single-parent families with incomes of less than €18,000 per year (in January 2019). Low-income pensioners remained by far the most numerous group of individuals affected by this policy (more than 80% of VC pensioners[2]).

The objective of this study was to estimate the impact of the VC intervention on the number of prescriptions dispensed. The study contributes in several ways to the literature on the demand for subsidized prescription medicines. First, our results show the effect of "zero-price" when a capped small coinsurance fee is replaced by free care, and they also compare this with the impact of a lower fixed co-payment (García-Gómez et al., 2018). Second, this is the first study that shows the effect of changes in pharmaceutical cost-sharing by constructing a "synthetic" counterfactual situation.

2 | DATA AND METHODS

2.1 | Data

We used a panel data from the reports published by the Spanish Ministry of Health and Social Policy[3] for 16 Spanish Autonomous Communities (ACs).[4] The data showed the volume of prescriptions in primary and specialized care dispensed from pharmacies and financed by the Spanish National Health System (SNHS). We used monthly data from July 2012 to December 2018, a period spanning months after the 2012 Spanish reforms, and after the abolition of charges for low-income pensioners, the intervention unit, that was implemented in the VC in January 2016. We computed the per-capita value as the total prescriptions dispensed divided by the resident population in each AC. Three predictor variables are proposed to account for the different behavior across ACs: The percentage of people over 65 years of age, the unemployment rate, and the average monthly air temperature[5] (degrees Celsius), see Table S1 for variable definitions and reference periods used.

The outcome variable used in this study, monthly per-capita total prescriptions, measures the number of prescriptions for total population while the intervention only affected low-income pensioners and disabled individuals given that only total data are available. However, there is evidence that the treated group represents a big share of the population and a bigger share of the total number of prescriptions. Average age of disabled individual in Spain was 64.3 years in 2007[6] and an important number benefit from a pension. In 2016, Spanish pensioners accounted for more than 8.6 million, representing 18.5% of total population, and at least 75%[7] of them are considered as low-income pensioners (≤18,000€). Data for 2017 show that out of the total number of prescriptions provided to pensioners, 75.4% corresponds to families with an annual income of less than €18,000.[8] Also, in 2017, 82.80% of the expenditure in medicines dispensed in pharmacies comes from pensioners and disabled individuals.[9] These figures show that low-income pensioners and disabled individuals represent a high proportion of the total number of prescriptions.

2.2 | Synthetic control method and Bayesian structural time series estimates

We performed our analysis using the synthetic control method (SCM), which first appeared in Abadie and Gardeazabal (2003) and became more formalized in Abadie, Diamond, and Hainmueller (2010) (hereinafter ADH model). In this approach, a weighted combination of potential control ACs, the synthetic control (SC), is constructed to approximate to the most relevant characteristics of the AC affected by the intervention (VC). After the intervention had been implemented in VC, it was possible to use SCM to estimate the counterfactual situation in the absence of intervention by looking at the outcome trend of the SC[10] (see the Supplementary Appendix for a detailed description of methods used for this study).

The recent literature on ADH SC proposes different methods to evaluate the predictive power of alternative sets of predictors. According to Dube and Zipperer (2016), we use a cross-validation criterion of minimizing the mean-squared prediction error among control units to choose between our candidate predictors (see Section S3 for more detail).

Following Abadie et al. (2010), we implemented the placebo tests based on permutation techniques; that is, we sequentially applied the ADH model to each AC in the group of potential controls and compared placebo with baseline results.

To the best of our knowledge, there is still no recognized way in the literature to generate confidence intervals for the SCM estimator when there is only one treated unit and a few control units. In order to provide an uncertainty estimate for the results, an alternative approach of SC estimation is proposed using Bayesian structural time series (BSTS; Scott & Varian, 2014). The main difference between the traditional SCM approach and the BSTS is that the latter models the outcome of the treated unit and uses the control units to capture the remaining variation. BSTS models aim to create an SC using the time series structure: trends and seasonality in prescribing before the start of the intervention, trends, and seasonality in other variables that are related to prescribing before the start of the intervention, and prior knowledge about model parameters that may be available from previous studies. By fitting the model to the pre-intervention period, the resulting parameter estimates can be used to predict counterfactual prescribing in the post-intervention period, based on the pre-intervention time series and post-intervention values for the covariates included in the model. The estimated impact of the intervention is then simply the difference between the observed and predicted values that would have occurred had the intervention not taken place (Brodersen, Gallusser, Koehler, Remy, & Scott, 2015).

A genuine Bayesian approach requires us to specify the prior distributions of all the model parameters. Section S4 describes the choice of prior distributions in the BSTS model.

We implemented our analysis in R using the *Synth* and *Causal Impact* packages (Abadie, Diamond, & Hainmueller, 2011; Brodersen, 2015).

3 | RESULTS

3.1 | Results using the synthetic control method

In light of the results of the cross-validation analysis, the ADH model includes four lags (the average value over July and August 2012, and the values in December 2012, April 2013, and March 2015) of the outcome variable in addition to one predictor variable: the percentage of population over 65 years of age. This predictor variable is biannual, averaged over the entire pre-intervention period.

Table 1 provides the numerical comparison for predictor variables and outcome variables between the VC and the constructed ADH SC. The SC accurately reproduced the values that the predictor variables had in the VC prior to the intervention. Table S2 shows summarized statistics for all data used in the empirical analysis, and Figure S1 plots the changes in per-capita prescriptions dispensed by ACs over time.

Comparisons of pre-intervention and post-intervention outcomes, in cumulative numbers, of the treated unit, and the SC provide the baseline estimates of the dynamic treatment effects. The cumulative result is a useful perspective if the response variable represents a flow quantity as in our case study. During the post-intervention period (January 2016 to December 2018), the per-capita prescriptions dispensed had a cumulative value of 64.39. In the absence of an intervention, we would have expected a cumulative per-capita prescriptions dispensed of 60.55. The cumulative per-capita prescriptions dispensed thus showed an increase of 6.34%.

The results from the ADH model are presented graphically in Figure 1. The average per-capita prescriptions dispensed over the months before the VC intervention was similar to that of the SC, root means squared prediction error (RMSPE) was equal to 0.0386, which included Asturias (25.6%), Cantabria (34.4%), Castile-La Mancha (26.5%), and Catalonia (13.5%; see Table 2). These weights indicate that per-capita prescriptions dispensed in the VC prior to intervention was best represented by a combination of these five ACs. All other ACs in the donor pool were assigned zero weights.

The results for the VC were strongly robust to placebo testing, as none of the placebo experiments for the 15-potential comparison ACs showed effects (a difference between each of the treated region's potential controls and their SC in the placebo study) larger than the baseline estimates. Figure 2 shows the result of the placebo study.

Finally, Table 3 displays a summary of significant effects during the 36 post-intervention months the vector of the number of placebo pseudo t-statistics (Galiani & Quistorff, 2017). In the first 30 months, the VC intervention significantly increased the per-capita prescriptions dispensed in the VC (15 for p-value < 0.05). However, for the last 6 months of our study period (from July to December 2018), a significant effect was only observed for 1 month for p-value < 0.05.

TABLE 1 Abadie et al. (2010) synthetic control model results: predictors and outcome means

	Valencian community	Synthetic	Difference (%)
Predictor variables			
Population over 65 years of age (%)	18.81	19.21	−2.08
The value of the outcome variable averaged over Jul and Aug 2012[a]	1.55	1.56	−0.64
The value of the outcome variable in Dec 2012[b]	1.62	1.51	7.28
The value of the outcome variable in Apr 2013[c]	1.62	1.64	−1.22
The value of the outcome variable in Mar 2015[d]	1.69	1.68	0.60
Outcome variable: per-capita prescriptions dispensed			
Pre-intervention cumulative per-capita prescriptions dispensed (Jul 2012–Dec 2015)	67.80	67.32	0.71
Post-intervention cumulative per-capita prescriptions dispensed (Jan 2016-Dec 2018)	64.39	60.55	6.34

Note: Intervention periods that better highlight the trend of the outcome before intervention.

[a]First months of the pre-intervention period.

[b]End of the first 6 months of the VC 2012 pharmaceutical coinsurance scheme reform.

[c]Four months after the increase in the monthly maximum pharmaceutical co-payment in pensioners.

[d]Three months after the modification of the co-payment on prescription drugs for civil servants.

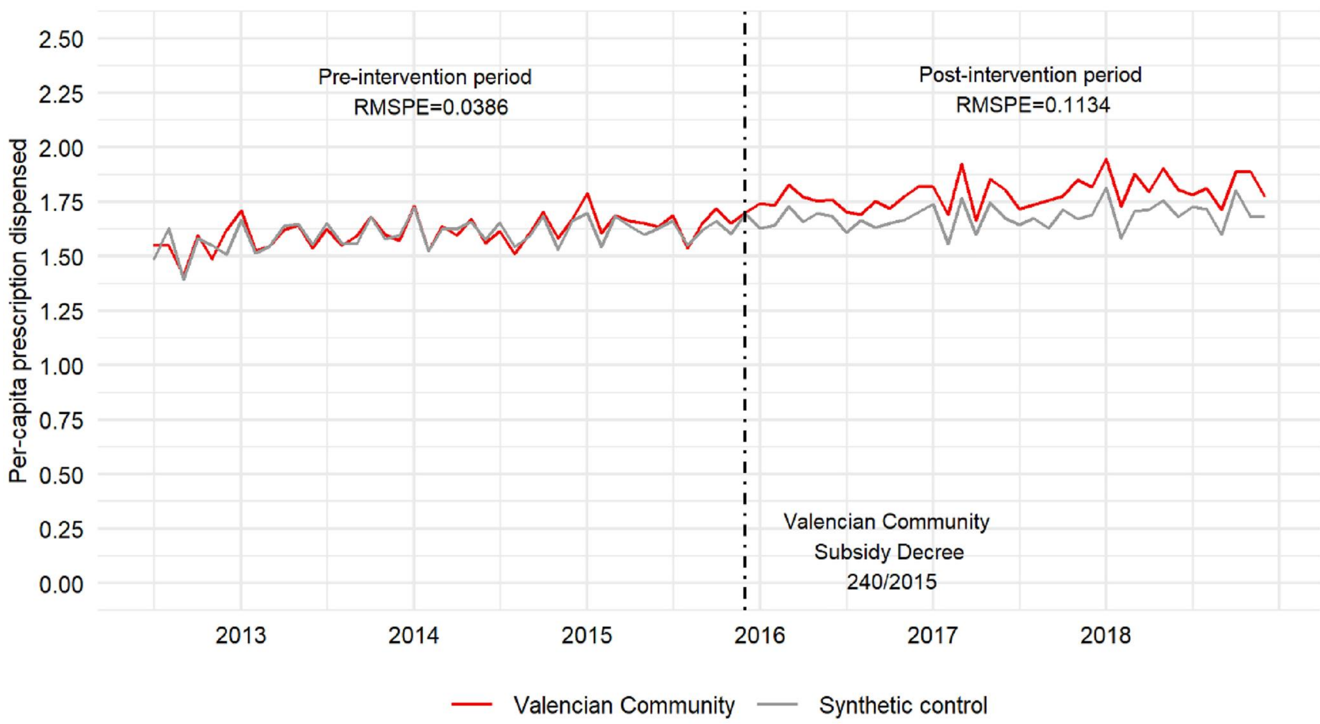

FIGURE 1 Trends in per-capita prescriptions dispensed: Valencian Community (VC) versus synthetic control (SC). RMSPE, root means squared prediction error

3.2 | Results using BSTS

Table 4 shows estimates obtained using the BSTS model. During the post-intervention period, monthly per-capita prescriptions dispensed in the VC had an average value of approximately 1.79. By contrast, in the absence of an intervention, we would have expected an average response of 1.67. The 95% credible interval (95% CI) of this

TABLE 2 Abadie et al. (2010) synthetic control model: estimated weights for each control unit

Autonomous Community (AC)	Weight
Andalusia	0
Aragon	0
Asturias	25.6%
Balearic Islands	0
Canary Islands	0
Cantabria	34.4%
Castile la Mancha	26.5%
Castile-Leon	0
Catalonia	13.5%
Extremadura	0
Galicia	0
La Rioja	0
Community of Madrid	0
Region of Murcia	0
Navarre	0

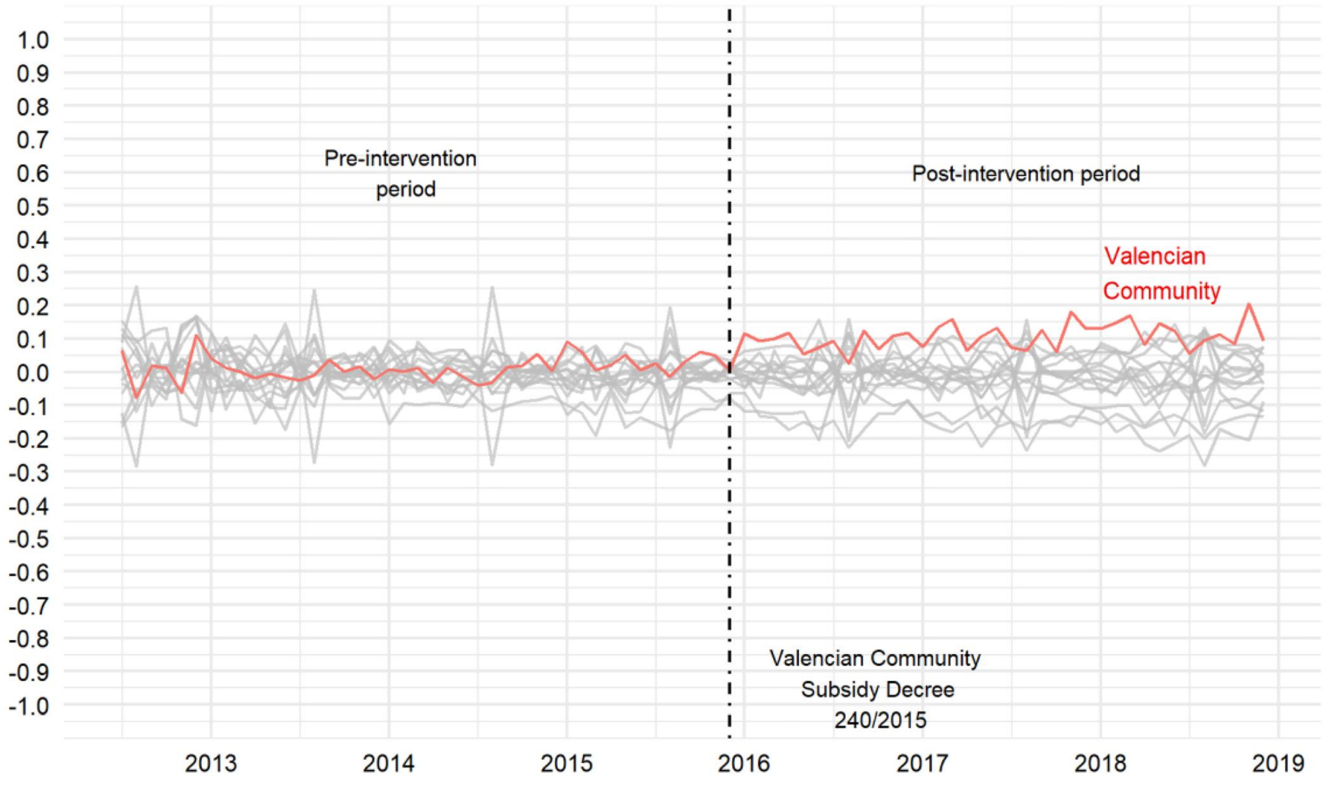

FIGURE 2 Abadie et al. (2010) synthetic control model: Difference in synthetic and observed per-capita prescriptions dispensed for Valencian Community and 15 placebo autonomous communities

counterfactual prediction is [1.63, 1.72]. Subtracting this prediction from the observed response yields an estimate of the causal effect which the intervention had on the response variable. This effect is 0.11 [95% CI: 0.07, 0.16]. Summing up the monthly data points during the post-intervention period, the response variable had an overall value of 64.39. By contrast, had the intervention not taken place, we would have expected a sum of 60.35 [95% CI: 58.66, 61.97].

TABLE 3 Abadie et al. (2010) synthetic control model: Summary of significant effects during the 36 post-intervention months

	Placebo test *p*-values (pseudo *t*-statistics)		
		Jan 2016 to Jun 2018	Jul 2018 to Dec 2018
		Number of months = 30	Number of months = 6
$p < 0.05$	Significant reduction	0	0
	Not significant	15	5
	Significant increase	15	1
$p < 0.10$	Significant reduction	0	0
	Not significant	12	5
	Significant increase	18	1

Notes: The *p*-values are calculated as the proportions of placebo effects that are at least as large as that of the treated unit (VC) for each post-intervention period. Then standardization, pseudo *t*-statistics, is achieved by dividing all *p*-values by the corresponding pre-treatment match root mean-squared prediction error in the pre-intervention period (Galiani & Quistorff, 2017).

TABLE 4 Bayesian synthetic structural time series model results

Response variable: per-capita prescriptions dispensed in Valencian Community		
Predictors: Trend and seasonal components and per-capita prescriptions dispensed in 15 remaining Autonomous Communities		
	Average	Cumulative
Actual	1.79	64.39
Prediction (s.d.)	1.67 (0.02)	60.34 (0.83)
95% CI	[1.63; 1.72]	[58.66; 61.97]
Absolute effect (s.d.)	0.11 (0.02)	4.04 (0.83)
95% CI	[0.07; 0.16]	[2.42; 5.73]
Relative effect (s.d.)	6.70% (1.38%)	6.70% (1.38%)
95% CI	[4.10; 9.47]	[4.10; 9.47]
Posterior tail-area probability *p*:	<0.0001	
Posterior probability of a causal effect	0.9999	

In relative terms, the response variable showed an increase of 6.70% [95% CI: 4.10, 9.47]. The probability of obtaining this effect by chance was very small (Bayesian one-sided tail-area probability $p = <0.0001$). BSTS results are graphically reported in Figure S2.

In the Supplementary Appendix, we include Table S6 summarizing the results of all different methods that have been conducted. We assessed the goodness-of-fit of each model by calculating the RMSPE between VC and synthetic ACs during the pre-treatment period. Comparing to differences in differences (DiD) model and SCM, BSTS decreased the RMSE a little bit to 0.0219, so we uphold this as our best model.

Also, in the Supplementary Appendix we include Table S7 with the average point effect obtained with BSTS for post-intervention divided in two subperiods. As suggested by the results presented in Figure S3, these results confirm that the 95% CI of the intervention impact is different from zero for the both period from January 2016 to June 2018 and from July to December 2018.

4 | DISCUSSION

This study used two novel SCM to estimate the impact on the number of dispensed prescriptions of a regional intervention which completely subsidized the national 10% pharmaceutical coinsurance rate, with a monthly ceiling of €8, until free access was granted for low-income pensioners and disabled individuals in the VC. Estimates using SCM and BSTS were very similar and showed a significant overall increase of 6.34% and of 6.70% [95% CI: 4.05, 9.47] in the number of prescriptions dispensed in this region, respectively. These figures correspond to an increase of about 11% in the number of prescriptions dispensed to those individuals affected by the intervention. Our results are in line with previous studies indicating that reducing price from a small amount to zero discontinuously boosts demand (Iizuka & Shigeoka, 2018; Puig-Junoy et al., 2016; Shampanier, Mazar, & Ariely, 2007).

These results provide empirical evidence that the impact of this intervention on the budget of the regional health service is far greater than the amount of the subsidy given the increase in the number of prescriptions. The overall incremental budget impact of the assessed intervention may be easily decomposed into: (i) the amount of the new subsidy introduced in 2016 in the VC assuming that the intervention does not have any impact on the number of prescriptions, (ii) the cost of the new subsidy corresponding to the increase in the number of prescriptions, and (iii) the cost for the public payer of this increase in prescriptions in addition to the new subsidy, which depend on the increase in prescriptions attributed to the intervention and the proportion of its price paid by the public payer since before implementing the intervention until today. Using macro data, we calculated that the overall impact on the budget may be an increase around 12% of public expenditure on outpatient medicines.

These results are useful for making accurate budgetary projections for similar eliminations of charges for low-income pensioners in the SNHS. The Canary Islands Autonomous Community (CI) enacted a similar subsidy for low-income pensioners from January 2019.[11] More importantly, the Spanish government led by the *Partido Socialista Obrero Español* (PSOE) has announced a similar elimination of the capped coinsurance rate for low-income pensioners,[12] despite the fact that non-pensioners with the same income continue to pay a noncapped 40% coinsurance rate.

In the VC and the CI, the rationale for this policy was a claimed effectiveness of the subsidy intervention in the form of an increase in treatment adherence for this population group.[13] Until now, only two studies have examined (with mixed evidence) the impact of the 2012 cost-sharing in Spain (Aznar-Lou et al., 2018; López-Valcárcel et al., 2017). López-Valcárcel et al. (2017) observed that the cost-sharing change made no significant differences in adherence among low/middle-income pensioners for low-priced essential medicines, such as antiplatelet and beta-blockers, but adherence temporarily declined for costlier Angiotensin-converting-enzyme inhibitor or an angiotensin II receptor blocker and statins. Aznar-Lou et al. (2018) reported a short-term increase in nonadherence to initial medication among low/middle-income pensioners after the introduction of the capped 10% coinsurance rate and a low regional co-payment in Catalonia. Then, a positive effect from the increased number of prescriptions associated with the suppression of the co-payment for pensioners and disabled individuals could have been a reduction in the risk of nonadherence to appropriate treatments. The reduction in nonadherence has been the continuously claimed objective of the intervention by the Valencian government.[14] This positive effect should be adequately confirmed by further research and it could not be simply inferred or denied with time series data used in this study. Conversely, this adherence increase effect will be lower when negative effects of the 2012 policy were only temporary and when the adherence reduction had been concentrated in the active population which faced a more intense co-payment than pensioners before and after 2012 policy.

Previous studies on the cost-sharing 2012 reform implemented in Spain in 2012 reported reduction in the number of prescriptions and daily defined doses around 13% one year latter (Antoñanzas et al., 2014; Hernández-Izquierdo et al., 2019) and ranging from 10% to 20% among ACs (Puig-Junoy et al., 2016). Our results for the suppression of the 2012 policy for low-income pensioners and disabled individuals in VC show a lower overall impact on the number of prescriptions (6.03%–6.69%). Notwithstanding, our results could not be easily compared to those obtained for the 2012 cost-sharing policy because this latter intervention included a cost-sharing increase for active population and non-low-income pensioners and, it also included a delisting of a long list of medicines. Despite that our method and data are not suitable to confirm or reject the hypothesis about the two policies being symmetric, the magnitude of the impact observed for the reverse policy implemented in VC in 2016 is lower than the impact of the 2012 policy.

Finally, it is important to improve the evidence on how long the effect of cost-sharing increases or decreases last for. Regarding the Spanish experience for the 2012 policy, two studies provided evidence suggesting that cost-sharing increases may have not a permanent effect on the reduction in the number of prescriptions and on adherence reduction (López-Valcárcel et al., 2017; Puig-Junoy et al., 2016). Our results provide evidence of a sustained significant effect of the VC intervention at least for 30 months after the intervention; however, we did not find significant effect for the months 31–36 after the intervention (Table S4). Also, the magnitude of our estimated impact for the first 30 post-intervention months cannot be fully interpreted as evidence of only a lasting effect given that the Valencian government marginally extended the subsidy to very small groups of population in 2017 and 2018 (to low-income children and adolescent until 18 years in 2017 and to low-income unemployed and irregular residents in 2018).

This study is affected by some limitations. First, our database did not allow us to estimate the impact of the intervention on adherence to treatment or on heterogeneity among therapeutic prescription groups. Second, our time series database did not allow us to estimate price elasticities or the effect of the intervention on the use of other healthcare services such as primary care or hospitalization (cross effects). Third, we were unable to appropriately verify symmetry of changes in opposite directions as it was the case in Spain in 2012 and in the VC in 2016. Fourth, more research is needed in order to get better evidence on the permanent or temporary effect of the analyzed intervention. Finally, a previous systematic review (Kill & Houlberg, 2014) of the effects of cost-sharing in health services identified several studies analyzing demand effects, and only a few of them analyzing distributional effects (equity), and health effects and substitution. Unfortunately, our data only allowed to estimate changes in consumption at the level of the number of prescriptions. Further research should consider a comprehensive analysis of welfare consequences should include all this effects, including heterogeneous effects and changes in adherence and appropriate use.

Future research should examine whether the increases in the number of prescriptions dispensed and in potential treatment adherence caused by the intervention in the VC reflect beneficial or low-value care (Baicker, Mullainathan, & Schwartzstein, 2015), and should examine cost-sharing alternatives that substantially reduce the reported “zero-price” effect.

CONFLICT OF INTERESTS

We don't have a conflict of interest to declare. We further attest that we have no commercial associations (e.g., equity ownership or interest, consultancy, patent and licensing agreement, or institutional and corporate associations) that constitute a conflict of interest in relation to the submitted manuscript. All sources of funding in support of the study are indicated in the appropriate section of the manuscript.

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ENDNOTES

[1] Source: http://www.san.gva.es/web/dgfps/aportacion-economica-1
[2] Source: https://w6.seg-social.es/ProsaInternetAnonimo/OnlineAccess?ARQ.SPM.ACTION=LOGIN&ARQ.SPM.APPTYPE=SERVICE&ARQ.IDAPP=ESTA0001. In the first half of 2018, 25% of the VC population had benefited from this subsidy, according to official sources.
[3] Source: http://www.mscbs.gob.es/profesionales/farmacia/datos/
[4] Because the 2012 Spanish reform was not applied in the Basque Country until the following year, this AC was excluded from the control group.
[5] There is a lot of evidence showing that there are external factors, including environmental conditions such as temperature and air quality, which affect certain pathologies that require the prescription of medicines.
[6] Source: www.ine.es
[7] Source: w6.seg-social.es.
[8] Source: https://www.msssi.gob.es/en/estadEstudios/estadisticas/estadisticas/estMinisterio/SIAP/home.htm
[9] Source: https://www.airef.es/en/spending-review.-study-2—medication-dispensed-through-prescription/
[10] In health economics, the SCM has been used, for example, to estimate the effect of pay-for-performance (Kreif et al., 2016; Ryan, Krinsky, Kontopantelis, & Doran, 2016), and universal health insurance coverage (Courtemanche & Zapata, 2014; Rieger, Wagner, Mebratie, Alemu, & Bedi, 2019). For more applications, consult the review by Bouttell, Craig, Lewsey, Robinson, and Popham (2018). Relative to DiD regression model-based studies, the SCM allows for substantially more flexibility in the treatment assignment mechanism, provided that a good balance in pre-treatment outcomes and observed covariates is achieved. SCM also does not require the strict assumptions for accurate estimation as in DiD methods. DiD method require average change in the nontreatment outcome to be the same among the controls and treated, making their paths parallel (Ryan, Burgess, & Dimick, 2015).
[11] Source: http://www.gobiernodecanarias.org/boc/2019/093/index.html
[12] Programa Electoral PSOE. Elecciones generales 2019. https://www.psoe.es/media-content/2019/04/PSOE-programa-electoral-elecciones-generales-28-de-abril-de-2019.pdf
[13] Source: https://www.micof.es/ver/17237/comunidad-valenciana-destaca-los-resultados-obtenidos-con-los-127-millones-gastados-en-ayudas-al-copago.html
[14] Source: http://www.san.gva.es/ayudas-copago

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SUPPORTING INFORMATION

Additional supporting information may be found online in the Supporting Information section at the end of this article.

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